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Welfare, Workfare and Labor Supply: A Unified *Ex Post* and *Ex Ante* Evaluation*

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Abstract

This paper analyzes the extent to which labor supply adjusts to incentives created by social programs. We find new evidence of highly elastic labor supply for single mothers in the United States, with sizable responses to the Earned Income Tax Credit (EITC) and welfare (AFDC/TANF) reforms during the 1990s. We reconcile some conflicting results in the literature by showing how the difference in differences design fails to identify a meaningful treatment parameter when a reform expands a pre-existing social program and when multiple programs change simultaneously. Finally, we use our quasi-experimental estimates to identify a structural model of labor supply with multiple tax and transfer programs. Model counterfactuals show that the effect of the EITC on labor supply depends on the regime of taxes and transfers in place. We conclude that evidence-based policymaking must explicitly model the tax and transfer system when using past reforms (*ex post* analysis) to draw inference about the effects of future reforms (*ex ante* analysis) on the labor market.

JEL Classification: H30, I38, J08, J38.

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1 Introduction

In fiscal year 2019, the United States spent $361 billion on income security programs. These social programs combine to generate a set of labor supply incentives. But do individuals respond to these incentives? Would labor supply differ if all social program money came in the form of conditional or unconditional cash transfers? The importance of this topic has generated a large body of literature that primarily conducts one of two types of analysis. The first type concerns the impact of past reforms enacted in a specific social, economic, and institutional context. We refer to this approach as *ex post* policy evaluation. The second type, *ex ante* policy evaluation, analyzes the counterfactual effects of policies that have never before been implemented.[1]

In this paper, we conduct *ex post* and *ex ante* evaluations of one of the largest set of recent policy reforms in the United States. Reforms to the Earned Income Tax Credit (EITC) and traditional welfare during the 1990s transformed the US social safety net from a welfare-oriented to a workfare-oriented regime. The EITC, a tax credit for parents who work and have low earnings, was progressively expanded to become the largest income support program in the US. Meanwhile a series of statewide reforms curtailed the generosity of welfare and culminated in the nationwide replacement of Aid for Families with Dependent Children (AFDC) with the stricter Temporary Aid for Needy Families (TANF).[2]

Central to any *ex post* analysis is the choice of method. We show that the difference in differences (DiD) design frequently used in the literature is inappropriate for the analysis of expansions of social programs with continuous levels of treatment. When a new program is introduced, DiD identifies a weighted average of traditional treatment parameters corresponding to each level of treatment, but when a program is merely *expanded* – as the EITC was in the 1990s – DiD does not identify a treatment parameter corresponding to the causal effect of the program. Even if a program has a positive effect on all individuals, DiD could

[1] Todd and Wolpin (2008) distinguish between *ex post* and *ex ante* analyses. Closely related are the discussion of Marschak's Maxim in Heckman (2001) and the taxonomy of policy evaluation problems in Heckman and Vytlacil (2007).

[2] Throughout the paper, we use the term welfare to refer to both AFDC and TANF.

produce an estimate that is zero or negative when the event being analyzed is an *expansion* of the program.

Our *ex post* empirical analysis aims to answer an economically meaningful question: Do individuals respond to the incentives generated by social program? The answer to this question lies in identifying the *marginal* responses of individuals to EITC and welfare benefits. To answer this question, we use data from the March Annual Social and Economic Supplement (ASEC) to the Current Population Survey (CPS) for the years 1988-2002, a period spanning the major reforms to welfare and the EITC during the 1990s. We examine single mothers as they have been shown to have elastic labor supply (Gelber and Mitchell 2012).

Our identification strategy relies on two main elements. First, the longitudinal nature of the data allows us to eliminate the effect of individual unobserved time-invariant heterogeneity by estimating the model in first differences. Second, we construct variables representing exogenous policy-induced changes in EITC and welfare benefits, similar to Dahl and Lochner (2012) and Agostinelli and Sorrenti (2018). Regressing changes in labor supply on changes in exogenous policy-induced benefits yields a consistent estimate of the average marginal response of labor supply during the period under investigation.

This *ex post* evaluation reveals that individuals are highly responsive to the incentives created by social programs. For each \$1,000 increase in EITC benefits, we estimate that labor supply increases by about 110 hours worked per year, an eight percent change relative to the pre-reform period. At the same time, a \$1,000 reduction in the generosity of welfare causes a statistically significant increase of 17 hours per year, a one percent change. These estimates are robust to the inclusion of controls for individual and household characteristics, state fixed effects, state unemployment levels, and the introduction of state-specific welfare reforms.

Given our findings that single mothers respond to incentives, we then conduct *ex ante* analysis of how they would respond to other proposed changes to the tax and transfer system, such as changes to the EITC and the introduction of

Universal Basic Income (UBI).[3] To do this, we specify a static model of labor supply for single mothers and use our *ex post* estimates to identify and validate the model. Our first *ex ante* evaluation decomposes the effect of reforms to the EITC and welfare in isolation and finds that employment among single mothers would have been 7.5 (5.5) percentage points lower had the EITC (welfare) not been reformed.

We then use the model to show how the elasticity of labor supply varies with the underlying tax and transfer system. We evaluate an EITC reform that expands both the federal income limit and the maximum federal benefits in 1996 while simultaneously varying features of the tax regime. We find that the average treatment effect (ATE) on employment depends on the effective marginal tax rate and the progressivity of the tax system. The same reform generates a sizable ATE when the tax system is characterized by high marginal tax rates or relatively high progressivity, but a small ATE when either are low. This result is important for two reasons: (i) It can reconcile comparisons of *ex post* evaluations of EITC expansions made in years with different tax codes (e.g. 1975, 1993, and 2009); and (ii) it highlights the importance of modeling the entire tax and transfer system when using past reforms to make predictions about the effects of future reforms on the labor market.

Finally, we analyze the consequences of replacing the current social programs with Universal Basic Income (UBI), a policy that has recently gained currency in the political debate (The Economist 2020). Our model suggests that UBI's effects on labor supply depend on which program is eliminated: Employment and hours worked decrease if UBI replaces the EITC but increase if UBI replaces welfare. These aggregate effects hide vast heterogeneity in individual responses, with hours worked responding non-monotonically depending on whether the individual was previously benefiting from the now-eliminated social program.

Overall, our study accomplishes three main tasks. First, we show that DiD does not identify meaningful treatment parameters when a reform expands the levels

[3] Heckman (1993) discusses the important distinction between descriptive labor supply functions and structural supply functions for "out-of-sample policy analysis concerning responses to tax and welfare programs."

of treatment for a continuous treatment variable. DiD has a role to play in program evaluation, but it is not appropriate for every setting. Second, we provide new evidence of the *marginal* effects of tax credits and welfare reforms on labor supply. Single mothers, it turns out, do respond to the incentives generated by social programs. Third, our *ex ante* analysis shows how the labor supply response to any particular reform depends on the entire set of incentives generated by the multiple social programs in place.[4]

Relationship to Literature. The literature that evaluates the effects of tax and transfer regimes on female labor supply is large, hence we limit our review to three strands that we believe are most closely related to ours: those works that evaluate the EITC using DiD designs, works that evaluate the EITC using a broader array of methods, and papers that explore the impacts of welfare on labor supply, fertility, and other outcomes.

Many *ex post* evaluations of the EITC implement various DiD designs that compare a control group of single women (who were ineligible for EITC benefits prior to 1994) to a treatment group of single mothers. Eissa and Liebman (1996) study the impact of the 1986 EITC reform via a DiD analysis, and find that the reform increased labor force participation for single mothers. With a similar empirical strategy, Meyer and Rosenbaum (2001) implement a decomposition analysis and find that a large share of the increase in employment by single mothers between 1984 and 1996 can be attributed to the EITC, with small effects due to welfare reform.[5] Kleven (2020) argues the opposite. He augments an event study of the 1993 EITC reform (one of the largest) with controls for state-specific welfare reforms and unemployment rates and shows that this shrinks the effect of the EITC

[4]For futher discussion about the limits of *ex post* evaluations without a theoretical framework that guides the understanding of underlying mechanisms, see Cunha and Heckman (2007), Heckman and Navarro (2007), Heckman and Urzúa (2010), Deaton (2010), Wolpin (2013), Heckman (2020) and Deaton (2020).

[5]These findings align with the results for a sample of California residents in Hotz, Mullin, and Scholz (2006), with studies based on longitudinal data such as Gelber and Mitchell (2012) and Agostinelli and Sorrenti (2018), and with studies exploiting event study setups around the largest EITC reforms such as Hoynes and Patel (2018). Dickert, Houser, and Scholz (1995), Meyer (2002), Eissa, Kleven, and Kreiner (2008), Bastian (2020), and Bastian and Lochner (2020) also find EITC-induced increases in maternal labor force participation and employment with the effect mainly driven by the group of single mothers. Blundell et al. (2016) show positive and large labor supply responses of single mothers to UK tax credits.

on the observed rise in employment in the 1990s to zero.

The discrepancy between the results obtained by our *ex post* empirical evaluation and that of Kleven (2020) can be explained by two factors. The first, as we have noted already, is that DiD designs do not identify the causal effect of the EITC on labor supply when applied to expansions of already-existing programs with continuous levels of treatment. We show that his approach of exploiting the differential introduction of welfare reforms across states further confounds the analysis without solving the identification problem. The second factor is the large degree of noncompliance in a design that treats all single mothers as "treated" and all single women as "controls." We show that over 30 percent of single mothers had no exposure to the EITC (due to high income) and that the analysis in Kleven (2020) is not robust to the exclusion of these untreated individuals from the "treatment" group.

The literature on the effects of the EITC on labor supply is summarized in excellent reviews such as Hotz and Scholz (2003), Eissa and Hoynes (2006), Meyer (2010), Nichols and Rothstein (2016), and Hoynes and Rothstein (2017). Hotz and Scholz (2003) discuss the importance of complementing reduced-form analyses with structural approaches that parameterize individual preferences and constraints in a theory of optimal decision making. Moffitt (1990), Keane and Moffitt (1998), and Keane (1995) build on this idea to analyze a wide range of policy reforms and expansions of the EITC. According to these studies, the EITC expansions between 1984 and 1996 considerably increased labor force participation, especially for the group of single mothers. Blundell et al. (2016), building on life-cycle models of female labor supply in, e.g., Heckman and Macurdy (1980), Eckstein and Wolpin (1989) and Keane and Wolpin (1997), show that the expansion of tax credits in the UK that are similar to the EITC increase single mothers' labor supply and marginally reduce educational attainment.[6]

Studies on the effect of welfare and welfare reforms on work incentives, welfare dependency, family structure, and migration are reviewed by Moffitt (1992), Blank (2002), and Grogger and Karoly (2005). Responses to changes in the welfare

[6]See also Blundell et al. (2000) for an *ex ante* analysis of the labor supply effects of a tax credit scheme in the UK before its actual implementation.

systems are multidimensional. For example, Moffitt (2019) analyzes the marginal treatment effects of AFDC reforms on labor supply. Moffitt, Phelan, and Winkler (2020) focus on changes in family structure. Keane and Wolpin (2007) analyze the effects of welfare on both labor force participation and fertility using a decision-theoretic life-cycle model of female labor supply. Grogger and Michalopoulos (2003) and Fang and Silverman (2004) show how the introduction of time limits to welfare recipients imply, under some conditions, lower welfare usage as well as improvements in the well-being of welfare recipients. Kline and Tartari (2016) find that the welfare reform experiment in Connecticut, Jobs First, induced sizable extensive margin responses for women. Meanwhile, the introduction of Food Stamps reduced employment and hours worked, especially for families headed by single mothers (Hoynes and Schanzenbach 2012).[7]

Empirically isolating the effect of changes in tax credits from separate changes in welfare is a central challenge in this literature. Ellwood (2000) argues that this task is particularly difficult, but finds that reforms to the EITC and welfare in the 1990s jointly caused an increase in the labor supply of single mothers. Grogger (2003) stresses that there are direct and indirect effects of the EITC on labor supply, with the indirect effect occurring through the EITC's causal reduction in individual welfare use. Fang and Keane (2004) show the substantial effect of the combined EITC and welfare reforms in shaping the increase in single mothers' labor force participation rate and the decrease in welfare caseloads.

Description of Programs. The EITC and welfare create contrasting incentives for labor supply behavior. To receive the EITC, a recipient must have a dependent child, positive earnings, and adjusted gross income below a threshold that varies with the year and number of dependent children.[8] The schedule of benefits depends on pre-tax income and features three parts: A phase-in where earned income receives a proportional subsidy, a plateau where benefits are neither increased nor reduced, and a phase-out where benefits are withdrawn. The incentives for recipients differ depending on their position in the schedule. Standard

[7] Mancino and Mullins (2020) find that EITC expansions generate positive responses of workers on transitioning into employment, transitioning to new jobs, as well as on accepting second jobs.

[8] Starting in 1994, individuals without dependent children but satisfying the other criteria were eligible for a small tax credit through the EITC.

models of labor supply predict that the substitution effect created by the phase-in would raise labor supply, while individuals situated on the plateau and phase-out would likely work less. The EITC's schedule of benefits was expanded continually throughout the 1990s, although one of the largest expansions was passed in 1993 and involved yearly increases in the benefit schedule for every year between 1994 and 1996.

In contrast to the EITC, traditional welfare has historically provided benefits to parents who do not work. Concerns that welfare disincentivized labor supply caused states to implement a variety of reforms, primarily between 1994 and 1996. These welfare waiver reforms contained a mix of incentives designed to encourage employment and reduce the number of families receiving benefits.[9] Many characteristics of the waiver reforms were eventually adopted nationwide when TANF replaced AFDC in 1996. Throughout this paper, we refer to the welfare-to-workfare transition as the set of welfare reforms and EITC expansions that were implemented in the United States between 1994 and 1996.

2 The *Ex Post* Evaluation and the DiD Estimand

Much of the empirical literature analyzing the effects of the EITC on labor market outcomes uses DiD or event study designs. A standard approach partitions women into single women without children, who are ineligible to receive EITC benefits prior to 1994, and single mothers, who *are* eligible, and compares changes in their labor supply around a reform using DiD (see for example Eissa and Liebman 1996; Meyer and Rosenbaum 2001; Kleven 2020; Schanzenbach and Strain 2020).[10]

In this section, we discuss what this standard DiD design identifies when the reform in question is not merely the introduction of a new program, but rather a reform of an already-existing social program. We discuss the identification results under different scenarios, emphasizing how the fuzzy DiD analyzed in the

[9] Appendix A contains a detailed description of the welfare waivers and changes to the EITC.

[10] Many studies consider single mothers as the treatment group, while single woman without children as the control group. This decision is based on the fact that single women were the primary targets of the 1993 EITC expansion, while women without children were excluded from receiving the EITC prior to 1994.

original work of De Chaisemartin and D'HaultfŒuille (2017) differs from our analysis of social program reforms.

Our framework considers a reform to a pre-existing social program that can change benefit levels in an unrestricted way. This feature is germane to any analysis of the EITC reforms that took place during the 1990s, which increased the income limits determining eligibility and the benefits of eligible individuals. Our focus on the standard DiD design is without loss of generality, as any year-specific coefficient in an event study can be estimated using a DiD design comprising only data from the year in question and the base year.

We consider the case where the set of possible EITC benefits is the discrete set $\mathcal{J} = \{0, 1, \ldots J\}$. Associated with each benefit level, j, at time t is a potential outcome $Y_{j,t}$. We define separate indicators, $D_{j,t}$, for each of the $J+1$ EITC benefit levels. At each point in time, the econometrician observes only one of the $J+1$ potential outcomes. We can write this observed outcome, Y_t, as a function of treatment assignments and potential outcomes:

$$Y_t = Y_{0,t} + \sum_{j=1}^{J} D_{j,t}(Y_{j,t} - Y_{0,t}) \; . \tag{1}$$

The standard DiD design estimates the following equation:

$$Y_{i,t} = \beta_0 + \beta_1 Post_t + \beta_2 Treat_i + \beta_3 Treat_i \times Post_t + \varepsilon_{i,t} \; , \tag{2}$$

where $Treat_i$ is a time-invariant binary variable equaling one if the individual receives EITC benefits at any point in the study, and $Post_t$ is a binary variable equaling one in the period after the reform.[11] The parameter of interest is β_3.

We consider a reform of the EITC that occurs between time $t-1$ and t. To facilitate discussion of this expansion, we introduce $D^*_{j,t}$, an indicator for the benefit level the individual would receive at time t if there had been no EITC reform, and a variable for the associated level of labor supply, $Y^*_{j,t}$. An expansion at time t that causes an individual to receive benefits equal to j instead of h would be

[11] Many single mothers are not eligible to receive the EITC because their incomes are too high (noncompliance). In Appendix B we explore the consequences of noncompliance.

characterized by a shift from $D^*_{h,t} = 1$ to $D_{j,t} = 1$. Because there is feedback between labor supply and EITC benefits – greater labor supply raises income, which in turn influences the benefit calculation – we interpret $D_{j,t}$ as an indicator variable for the benefits received by the individual as a result of the EITC reform only, before the feedback between labor supply and EITC benefit receipt has been taken into account. In Section 3 we discuss the empirical counterpart of the latent variables described above, while in Appendix B we analyze the case of imperfect compliance.

The first assumption we invoke is standard in the program evaluation literature that uses DiD designs:

Assumption 1 (No Selection on Counterfactual Trends).

$$E[Y_{0,t} - Y_{0,t-1}|Treat] = E[Y_{0,t} - Y_{0,t-1}]$$

Under Assumption 1, the DiD estimand represents the differential longitudinal changes in the outcome Y_t between treatment and control group: $\beta_3^{DiD} = E[Y_t - Y_{t-1}|Treat = 1] - E[Y_t - Y_{t-1}|Treat = 0]$. Below, we define two main causal parameters of interest.

Definition 1 (The EITC Target Parameters). *Two main parameters of interest with respect to an ex post evaluation of the EITC program are the average effect of treatment on the treated (ATT), and the average treatment effect of the policy reform (ATPR). These parameters are defined as follows:*

$$ATT \equiv \frac{1}{\sum_{j=1}^{J} \mathbb{P}(D_{j,t} = 1)} \sum_{j=1}^{J} \mathbb{P}(D_{j,t} = 1) E[Y_{j,t} - Y_{0,t}|D_{j,t} = 1] \ ,$$
$$ATPR \equiv \frac{1}{\sum_{j=1}^{J} \sum_{h=1}^{J} \phi^*_{j,h}} \sum_{j=1}^{J} \sum_{h=1}^{J} \phi^*_{j,h} E[Y_{j,t} - Y^*_{h,t}|D_{j,t} = 1, D^*_{h,t} = 1] \ ,$$

where $\phi^*_{j,h} = \mathbb{P}(D_{j,t} = 1, D^*_{h,t} = 1)$.

ATT is a causal parameter that measures the effectiveness of the program as it is currently implemented. It identifies that program's average effect across all levels of treatment relative to a world with no program *for the people who actually participate in the program.* While ATT evaluates the program itself, ATPR can be used to evaluate a *reform* of the program. It measures the effect of the program's reform for people whose benefit level was altered by the reform.

However, Proposition 1 clarifies that without additional assumptions, the DiD estimand does not identify either of the target parameters.

Proposition 1. *Suppose Assumption 1 holds. Then, the DiD estimand is proportional to the difference of weighted sums of treatment on the treated (TT) parameters for each treatment level:*

$$\beta_3^{DiD} = \frac{1}{p_{Treat}} \sum_{j=1}^{J} (p_{j,t}\Delta_{j,t}^{TT} - p_{j,t-1}\Delta_{j,t-1}^{TT}) \ , \qquad (3)$$

where $p_{Treat} = \mathbb{P}(Treat = 1)$*, and for each* $j = 1\ldots, J$*,* $p_{j,t} = \mathbb{P}(D_{j,t} = 1)$*,* $\Delta_{j,t}^{TT} = E[Y_{j,t} - Y_{0,t}|D_{j,t} = 1]$*, and* $p_{j,t-1}$ *and* $\Delta_{j,t-1}^{TT}$ *are defined analogously.*

Proof. See Appendix B. □

Proposition 1 reveals that the DiD estimand could be zero or negative even if the treatment effect at each treatment level, $Y_{j,t} - Y_{0,t}$, is positive and hence $\left\{\Delta_{j,t}^{TT} > 0\right\}_{j=1}^{J}$. Moreover, without additional restrictions, the DiD estimand in Equation (3) is silent about the average effect of the policy reform (ATPR). We now consider several special cases of the model and ask whether DiD identifies any causal parameters of interest. Appendix B considers additional cases including imperfect compliance. The Appendix shows that strong restrictions on the behavior of treatment and control individuals are required for the DiD estimand to have a causal interpretation in the presence of imperfect compliance.

No pre-existing policy regime. If a program is introduced at time t for the first time, so that no individuals in the treatment group were exposed to the program prior to the reform ($p_{j,t-1} = 0$ for all $j \neq 0$), DiD identifies ATT, a weighted average of treatment on the treated: $\beta_3^{DiD} = \sum_{j=1}^{J} \omega_{j,t}\Delta_{j,t}^{TT}$. The weights $\omega_{j,t} = \frac{p_{j,t}}{\sum_{j=1}^{J} p_{j,t}}$ are given by the fraction of the treated population receiving each level of the treatment, and they sum to one. However, this case is not relevant when the

DiD design is used to evaluate programs like the EITC that were already in place prior to the reform being analyzed.

DiD Interpretation with Additional Restrictions. We next consider the case in which the potential outcomes are time-invariant, $Y_{j,t} = Y_{j,t-1} = Y_j \; \forall j \in \mathcal{J}$. In this case, the DiD estimand is a function of the marginal effects on the treated generated by the reform: $\beta_3^{DiD} = \frac{1}{p_{Treat}} \sum_{j=1}^{J} \sum_{h \neq j} \phi_{j,h} E[Y_j - Y_h | D_{j,t} = 1, D_{h,t-1} = 1]$, where $p_{Treat} = \mathbb{P}(Treat = 1)$, and $\phi_{j,h} = \mathbb{P}(D_{j,t} = 1, D_{h,t-1} = 1)$.

The DiD estimand in this case identifies the ATPR only in a very special case. Identification of ATPR requires both $D^*_{h,t} = D_{h,t-1}$ for $h = 1, \ldots, J$ and for $\sum_{j=1}^{J} \sum_{h \neq j} \phi_{j,h} = p_{Treat}$. The first assumption means that individuals must have the same treatment level in time $t-1$ as they would have in time t absent the reform, while the second means that every individual must change the amount of EITC benefits they receive between periods $t-1$ and t because of the reform. These assumptions are unlikely to hold true in many empirical applications in the EITC literature: In Section 3.3 we show that the fraction of single mothers with no change in EITC benefits over time is sizable. Finally, restricting the potential outcomes over time can be inconsistent with changes in the rest of the tax and transfer programs, where an individual's labor supply for a given level of EITC benefits can change because of incentives from other social program reforms.

Similar restrictions were made in De Chaisemartin and D'HaultfŒuille (2017), who analyze the identification of treatment effects in a fuzzy DiD design. Although we acknowledge the relevance of their results, we also believe that the proposed estimands are unlikely to be applicable in the case of social welfare reforms. First, both single mothers and single women who receive no EITC benefits before the reform are likely to differ in their subsequent labor supply after the reform, violating their conditional common trend assumption. This can be driven by differences in preferences, as well as by other changes in the tax and transfer system, such as the TANF reform, that affect single mothers but not single women. Moreover, De Chaisemartin and D'HaultfŒuille (2017) consider the average causal response (ACR) as the targeted causal parameter of the proposed estimands. This causal parameter is not well defined in our framework, as social welfare reforms like the EITC reform typically cause more than a one-unit increment in the treatment intensity, generating potential overlapping of complier groups among different margins of the treatment (see Angrist and Imbens 1995).

Simultaneous changes in other tax and transfer programs. As noted in Kleven (2020), the effect of the EITC on individual behavior is unlikely to be constant over time because of changes in the rest of the tax and transfer system. This makes the above restrictions that $Y_{j,t} = Y_{j,t-1} = Y_j$ for $j \in \mathcal{J}$ unrealistic. The effect of a tax credit on labor supply before welfare reform may differ from the effect of the same credit after welfare reform.

Kleven (2020) discusses how state-level reforms to AFDC programs occurred contemporaneously with the 1993 EITC expansion, and the DiD design confounds the effects of the two policy reforms on employment. The author proposes overcoming the identification challenges posed by simultaneous reforms by exploiting the uneven introduction of welfare waiver reforms across states. In this framework, $W_{i,t} \in \{0,1\}$ indicates whether individual i resides in a state that had implemented welfare waiver reforms ($W_{i,t} = 1$) or not ($W_{i,t} = 0$) by time t. The redefined DiD model including the effects of waivers is

$$\begin{aligned} Y_{i,t} =& \beta_0 + \beta_1 Post_t + \beta_2 Treat_i + \beta_3 Post_t \times Treat_i + \\ & \beta_4 W_{i,t} + \beta_5 W_{i,t} \times Post_t + \beta_6 W_{i,t} \times Treat_i + \beta_7 W_{i,t} \times Post_t \times Treat_i + \varepsilon_{i,t} \, . \end{aligned} \tag{4}$$

This approach compares employment rates in states that implemented welfare waivers to those that did not in an effort to estimate a treatment parameter for the EITC. However, as Proposition 2 makes clear, this approach still does not estimate a meaningful treatment parameter.

Proposition 2. *Suppose assumption 1 holds. Then the DiD estimand is:*

$$\beta_3^{DiD} = \frac{1}{p_{Treat}} \sum_{j=1}^{J} p_{j,t}^{W=0} \Delta_{j,t}^{TT,W=0} - p_{j,t-1}^{W=0} \Delta_{j,t-1}^{TT,W=0} \, , \tag{5}$$

where $p_{Treat} = \mathbb{P}(Treat = 1 | W_{i,t} = 0)$*, and for each* $j = 1 \ldots, J$*,* $p_{j,t}^{W=0} = \mathbb{P}(D_{j,t} = 1 | W_{i,t} = 0)$*,* $\Delta_{j,t}^{W=0} = \mathbb{E}(Y_{j,t} - Y_{0,t} | W_{i,t} = 0)$*,* $p_{j,t-1}^{W=0} = \mathbb{P}(D_{j,t-1} = 1 | W_{i,t} = 0)$*, and* $\Delta_{j,t-1}^{W=0} = \mathbb{E}(Y_{j,t-1} - Y_{0,t-1} | W_{i,t} = 0)$.

Proof. See Appendix B. □

Proposition 2 reveals that the DiD estimand when controlling for waivers is still a difference in weighted treatment on the treated parameters for each margin of the EITC program. Moreoever, the use of event study designs when controlling for waivers introduces an additional complication. Each coefficient in the event study is identified by conditioning on the subset of states that had not implemented any welfare waivers by the time of the year in question. As the set of states that had not implemented welfare waiver reforms by 1994 will differ from the set of states not having implemented them by 1995, the coefficients for 1994 and 1995 in the event study will be identified by labor supply changes in a different set of states. This means that comparisons of the event study coefficients across years will confound changes in employment over time resulting from the reform with changes in the average employment level that result from conditioning on a different set of states.[12] Therefore, not only do the event study coefficients fail to identify any meaningful target parameter related to the 1993 EITC expansion, comparisons across years are also misleading, because the set of states identifying each coefficient varies by year.

Kleven (2020) concludes that the negligible effect of the EITC on labor supply – after controlling for the waiver states – is a result of informational or pyschological frictions. We propose an alternative explanation: The DiD estimator is not well-suited to the *ex post* evaluation of reforms of pre-existing social programs. We next propose a different empirical model that can identify a meaningful parameter: the average marginal effect of the EITC on labor supply. We will also identify the same parameter for welfare. The model will reveal that the EITC and welfare each have large effects on the labor supply of single mothers.

3 New *Ex Post* Evaluation of the EITC and Welfare Reforms

In this section, we estimate causal effects of expansions of both the EITC and welfare on the labor supply of single mothers. We focus on the marginal effects of

[12] Grogger and Karoly (2005) discuss the timing of the implementation of the welfare waiver reforms.

each policy, which we use to determine the extent to which individuals respond to the incentives created by these programs.

Our main empirical specification of labor supply for individual i at time t is

$$Y_{i,t} = \beta_0 + \gamma_0 t + \gamma_1 \xi_{i,t} + \gamma_2 T_{i,t} + \alpha_i + \epsilon_{i,t} \ , \tag{6}$$

where $Y_{i,t}$ is a measure of labor supply, $\xi_{i,t}$ is the real value of EITC benefits, $T_{i,t}$ is the real value of welfare benefits, α_i represents an individual-specific unobserved preference for work, and $\epsilon_{i,t}$ represents additional unobserved heterogeneity. We allow for a possible time trend, denoted by t. The marginal effect of EITC benefits is given by γ_1, while the marginal effect of welfare benefits is γ_2.

Before estimating (6), we difference the equation to eliminate each individual's unobserved preference component, α_i, yielding

$$\Delta Y_{i,t} = \gamma_0 + \gamma_1 \Delta\xi_{i,t} + \gamma_2 \Delta T_{i,t} + \Delta\epsilon_{i,t} \ . \tag{7}$$

Following this transformation, the average marginal effects of EITC and welfare benefits are identified under the following standard exogeneity assumption.

Assumption 2. *The longitudinal change in the unobserved heterogeneity of labor supply among individuals is mean independent with respect to the policy-induced longitudinal changes in EITC and welfare benefits:* $E[\Delta\epsilon_{i,t} \,|\, \Delta\xi_{i,t}, \Delta T_{i,t}] = E[\Delta\epsilon_{i,t}] = 0$.

Assumption 2 is similar to the standard parallel trend assumption in DiD, although here the unobserved determinants of labor supply, $\Delta\varepsilon_{i,t}$, must be mean independent of the intensity of the treatment ($\Delta\xi_{i,t}, \Delta T_{i,t}$) rather than whether an individual is in the treatment group. Section 3.1 describes the way we construct $\Delta\xi_{i,t}$ and $\Delta T_{i,t}$ so that Assumption 2 is likely to be satisfied.

Under Assumption 2, Ordinary Least Squares (OLS) identifies γ_1 and γ_2. If labor supply is not truly linear in EITC and welfare benefits, then OLS identifies the *average* marginal effect of EITC benefits (welfare benefits) on hours worked, *ceteris paribus* for the level of welfare benefits (EITC benefits):

$$\gamma_1 = \frac{\partial E[\,Y_{i,t} \,|\, \xi, T]}{\partial \xi} \quad \text{and} \quad \gamma_2 = \frac{\partial E[\,Y_{i,t} \,|\, \xi, T]}{\partial T} \ .$$

We assess the linearity assumption in Section 3.2 and estimate a model allowing for nonlinear response in Section 4.

Unlike DiD designs, consistent estimation of the specification in (7) actually identifies an interesting treatment parameter: the average marginal effect of the treatment in the population. A nonzero marginal effect of the EITC would reveal whether single mothers respond to the incentives created by the program. Furthermore, if the marginal effects are constant, $Y_{j,t} - Y_{j-1,t} = \gamma$ for $j = 1, \ldots J$, they can be used to construct a back-of-the-envelope estimate of ATT by multiplying the marginal effect by the average EITC benefit among recipients and rescaling by the proportion of treated individuals in the population.

3.1 Data

We use the March Annual Social and Economic Supplement (ASEC) to the Current Population Survey (CPS). The CPS is a monthly US household survey that asks each member of the household detailed questions related to labor force participation, earnings, and demographic characteristics. Households selected into the CPS are surveyed eight times over a period of 16 months.

We use the ASEC supplement for several reasons. As the survey occurs once per year, we can create a panel with two observations per individual. Moreover, the ASEC measures labor force participation and income in the prior calendar year while the CPS monthly files measure variables in the prior week or prior twelve months. Since the EITC is calculated on the basis of annual earnings and our analysis relies on a precise continuous measure for policy-induced changes in EITC benefits over time, we view the ASEC as more appropriate for our analysis.

Our sample consists of single mothers between the ages of 25 and 50 who are present in two consecutive years of the ASEC supplement. We look at single mothers for three reasons: (i) women with children are a main target of programs such as the EITC and welfare; (ii) single mothers are deemed one of the most responsive groups to tax and welfare reforms; and (iii) it simplifies the analysis by abstracting from potential interaction effects with partners' behavior (Gelber and Mitchell 2012).[13]

We create longitudinal linkages for single mothers in the ASEC between 1988 and 2002, resulting in a sample of 10,959 unique mothers.[14] Each single mother is observed one time per year for two consecutive years. For each mother, we construct longitudinal (year-on-year) changes in total yearly hours worked.[15]

Following Dahl and Lochner (2012) and Agostinelli and Sorrenti (2018), we construct exogenous policy-induced changes in EITC and welfare benefits, $\Delta\xi_{i,t}$ and $\Delta T_{i,t}$, for each individual in the sample. This involves predicting the counterfactual level of earnings ($E_{i,t}$) and non-labor income ($NL_{i,t}$) that would have prevailed in the second period (t) of the sample in the absence of any labor supply response to the EITC and welfare reforms. $\widehat{E}_{i,t}$ (respectively $\widehat{NL}_{i,t}$) is the predicted value from a regression of $E_{i,t}$ ($NL_{i,t}$) on a fifth-order polynomial in its lag and an indicator for positive lagged values. $\widehat{E}_{i,t}$ and $\widehat{NL}_{i,t}$ then serve as inputs in the computation of second-period welfare and EITC benefits, ($\widehat{\xi}_{i,t}$,$\widehat{T}_{i,t}$), as follows:

$$\widehat{\xi}_{i,t} = EITC_{i,t}(\widehat{E}_{i,t}) \ ,$$
$$\widehat{T}_{i,t} = T^{Welfare}_{i,t}(\widehat{E}_{i,t}, \widehat{NL}_{i,t}) \, ,$$

where $EITC_{i,t}(\cdot)$ and $T^{Welfare}_{i,t}(\cdot)$ are functions that calculate EITC and AFCD/TANF benefits in year t based on program rules and individual characteristics (number of dependent children and state of residence). Our measures of exogenous policy-induced changes in benefits are then given by:

$$\Delta\xi_{i,t} = \widehat{\xi}_{i,t} - \xi_{i,t-1} \ , \quad (8)$$
$$\Delta T_{i,t} = \widehat{T}_{i,t} - T_{i,t-1} \ . \quad (9)$$

EITC benefits are calculated using NBER's TAXSIM, and welfare benefits are

[13]Prior to 1994, EITC recipients must report at least a dependent child to be eligible for the tax credit. Following the 1993 EITC reform, a small number of women without children were allowed to receive the EITC.

[14]Rivera Drew, Flood, and Warren (2014) discuss the challenges in linking individuals across years of the CPS due to sample attrition and individuals not matching on demographic characteristics in successive years.

[15]We construct the variable for the total hours worked per year by multiplying the total weeks worked per year by the ASEC variable denoting "usual hours worked per week."

computed from the AFDC/TANF rules in place for each year and state. Appendix D provides further details about the construction of these variables. Descriptive statistics for our sample are presented in Appendix Table C-1.

3.2 Empirical Results

In this section we provide estimates of the marginal effects of both the EITC and welfare on female labor supply. We also provide evidence of an approximately linear relationship between labor supply and benefits. At the end, we use our estimates of marginal effects to provide an estimate of ATT for each program that is valid under an assumption of linearity.

Figure 1 depicts the binscatter of the nonparametric relationship between the longitudinal change in hours worked by single mothers and the variables $\Delta\xi_{i,t}$ and $\Delta T_{i,t}$.[16] All specifications include controls for state unemployment level, state and year fixed effects, state welfare waivers, and indicator variables for race and the number of dependent children. The analysis of the effect of EITC changes also controls for changes in welfare benefits, and viceversa.

Two results are worth highlighting. First, Figure 1-a suggests a positive relation between workfare policies and single mothers' labor supply. Expansions in EITC benefits induce increases in the amount of hours worked by single mothers. Conversely, Figure 1-b highlights a negative relationship between changes in welfare benefits and maternal labor supply. A policy regime granting decreasing amounts of welfare benefits to single mothers generates increasing levels of labor supply. Second, the relations between EITC and welfare benefits and hours worked are well-approximated by a linear specification.

Table 1 presents the main result of our *ex post* analysis. In column (1), we estimate a model only including longitudinal changes in EITC benefits. In column (2), we focus on changes in welfare benefits in isolation. Column (3) includes both changes in EITC and welfare benefits. Column (4) augments the model in (3) with controls for state and year fixed effects, and includes indicator variables

[16] Figure C-2 in Appendix C replicates the analysis with changes in the logarithm of hours, thereby conditioning on the subsample working positive hours in both periods. The figure reveals a positive (negative) relationship between the EITC (welfare) and the intensive margin.

for race and the number of dependent children. Column (5) adds controls for state unemployment level and an indicator variable for state welfare waivers. The last model in column (6) further augments the specification with a control function in earnings and non-labor income to take into account the initial position of each mother in the EITC and welfare schedule. This approach is similar to the one proposed in Dahl and Lochner (2012) and Agostinelli and Sorrenti (2018). The control function includes lagged variables for labor income, business income, farm income, and non-labor income.

According to the specification in column (1) of Table 1, a $1,000 increase in EITC benefits induces a statistically significant increase of approximately 125 hours worked per year. The effect size is slightly lower, about 110 additional hours per year, and remains positive and highly significant in the specifications with additional controls in columns (2) to (6). The effect is of similar magnitude to the effect in Agostinelli and Sorrenti (2018).[17] Welfare benefits have the opposite effect. A $1,000 decrease in welfare benefits induces a statistically significant increase of about 32 hours worked per year in the specification with welfare benefits in isolation and by about 12-17 hours in more saturated specifications.[18]

The analysis of hours worked highlights two important insights. First, it confirms that workfare-oriented policy regimes cause a positive labor supply response among single mothers, while welfare-oriented regimes have the opposite effect. This means that the combination of an increase in the EITC and a reduction in welfare – namely, a welfare-to-workfare transition – causes an unambiguous increase in the aggregate labor supply of single mothers. Second, estimates for the effect of changes in both EITC and welfare benefits are robust to the inclusion of a wide range of controls for individual characteristics, such as race and number of children, and macroeconomic factors like state unemployment levels and the presence of state welfare waivers.

Table 2 investigates heterogeneity in labor supply responses to the EITC and wel-

[17] Agostinelli and Sorrenti (2018) estimate a similar specification for labor supply responses to EITC expansions over time with a different data set (NLSY79) and a sample consisting of both single and married (or cohabiting) mothers.

[18] Appendix Tables C-2 and C-4 extend the analysis to weeks worked and log-hours worked.

fare by employment status.[19] The results show that the increase in hours worked by single mothers is mainly driven by the expansion of the EITC program over time. We estimate an average of 70 extra hours of work per year in response to a $1000 increase in EITC benefits among single mothers who were not previously working. Conversely, this group is relatively unaffected by welfare, with small and statistically insignificant point estimates in the most complete specifications. Among working single mothers, each $1000 decrease in welfare benefits causes an increase in labor supply of about 26 hours per year. The impact of the EITC expansion for the subgroup of working single mothers is of the same magnitude but statistically insignificant.[20]

These results align closely with the incentives created by the two programs. For mothers already working, an EITC expansion has a negative effect on labor supply for individuals located on the plateau or downward-sloping part of the schedule, which partly offsets the strong substitution effect created by the schedule's upward-sloping segment. The labor supply incentives of an EITC expansion are, however, unambiguously positive for nonworking mothers, which is reflected in the much larger and more precise point estimate. Likewise, reductions in welfare remove the incentive to have low earnings among those already working, which explains the negative and statistically significant estimates in Table 2.

We can use the point estimates of marginal effects from column (6) of Table 1 to construct a back-of-the envelope estimate of ATT for each policy. When there is no heterogeneity in labor supply and when the marginal effects are constant, ATT is simply the product of the marginal effect and the average EITC benefit among recipients rescaled by the proportion of treated individuals. Averaged over all of the years in our sample, this back-of-the-envelope calculation yields an ATT 309 hours per year for the EITC, while for welfare it is -262 hours per year. In computing ATT, the much smaller marginal effects for welfare relative to the EITC are offset by the larger average benefit level among single mothers during the sample period. The existence of heterogeneity in labor supply for a fixed benefit profile and as well as a nonlinear labor supply function would bias

[19] Employment status is measured in the first ($t-1$) of the two time observations we have for each mother in the sample.

[20] The analysis of weeks worked in Appendix Table C-3 suggests similar insights.

these calculations. We allow for these departures in the model in Section 4.

3.3 Reconciling Previous Literature with Our Results

Much of the empirical literature on the effect of expansions in the EITC on labor supply points to positive extensive margin responses, especially for the group of single mothers (e.g. Eissa and Liebman 1996 and Meyer 2002). On the other hand, recent analysis by Kleven (2020) shows no effect of the EITC on single mothers' labor supply. In this section, we explore how heterogeneity in benefits within the treatment group drives labor supply responses and how this is missed by the standard DiD design.

Figure 2 shows that there is substantial heterogeneity in the benefits that are induced by reforms to both the EITC and welfare. Figure 2-a shows that policy-induced changes in EITC benefits are large for the poorest single mothers, slightly negative for mothers earning between \$12,000 and \$22,000, and close to zero for everyone else.[21] The figure also demonstrates heterogeneity in changes in labor supply. The left tail of the earnings distribution has the largest increase in hours worked (up to 300 additional hours per year), while mothers earnings between \$10,000 and \$30,000 tend to reduce their hours. Mothers with still higher earnings experience no change.

Figure 2-b repeats the exercise for changes in welfare benefits. There is a strong negative relationship between changes in hours and welfare benefits only among individuals with low incomes (below \$10,000), suggesting that most of the identifying variation for the effect of welfare on hours comes from the poorest single mothers, an observation consistent with welfare's eligibility rules. Together, Figures 2-a and 2-b demonstrate substantial heterogeneity in benefits among single mothers.[22]

The standard DiD/event study design used in the literature reduces all this heterogeneity within the treatment variable to a single margin, that of going from

[21] Our earnings prediction model typically predicts higher earnings in the second period (t) relative to the first period ($t-1$), which can result in lower EITC benefits for mothers located on the plateau or phase-out of the EITC schedule as the EITC is expanded.

[22] Fang and Keane (2004), among others, highlight the existence of heterogeneous effects of welfare by showing that about one-quarter of welfare leavers did not start working.

untreated to treated, and eliminates the relationship between the intensity of treatment and the intensity of the labor supply response. Furthermore, while single mothers form the treatment group, a large fraction of them were never exposed to the policy reform. Figure 3 plots the distribution of policy-induced changes in EITC benefits due to changes in the program schedule for single mothers. Two patterns are striking. First, changes in EITC benefits display a high level of variation. Second, more than one third of single mothers are unaffected by expansions of the EITC during the period of consideration for the event study.[23] The practice of assigning unexposed mothers to the treatment group confounds the resulting estimate in much the same way that imperfect compliance in a randomized controlled trial (RCT) affects the interpretation of the RCT evaluation as ITT instead of ATE.

We therefore investigate the robustness of the research design in Kleven (2020) to the exclusion of single mothers who received no benefits from the treatment group. If the labor supply of single mothers were unaffected by the EITC expansion, the event study would be robust to this exclusion.[24]

Figure 4 displays these event studies of the 1993 EITC expansion.[25] The results are obtained by regressing each outcome on the interactions between the indicator variable for treatment and indicator variables for each year, after controlling for state waivers and unemployment levels.

Figure 4-a shows that in the post-reform period, the treatment effect of being a single mother is positive (135 additional hours per year) and statistically significant at the ten percent level in 1995 and at the five percent level in 1996. Figure 4-b replicates the analysis on the extensive margin done in Kleven (2020). The fig-

[23]Many single mothers are unexposed by changes in the EITC schedule because their earnings are too high to be eligible. No labor supply response to the EITC expansion should be expected for these mothers.

[24]To perform the event study, we have augmented the sample by including single women without children, a group that was not part of our empirical analysis in Section 3.2. Single women without children form the control group since, prior to 1993, the presence of at least one dependent child was a requirement for EITC eligibility.

[25]The only difference between the analysis here and that of Kleven (2020) is our use of the ASEC instead of the full CPS. Our analysis of the extensive margin in Figure 4-b produces point estimates similar to those in Kleven's study, but with larger standard errors due to the smaller sample size.

ure provides the same qualitative conclusions as the analysis of hours worked, but the point estimates are noisy, and we cannot reject the hypothesis that the 1993 EITC reform had no impact on the extensive margin.

Figures 4-c and 4-d replicate the above analysis excluding single mothers with no change in EITC benefits according to the variable plotted in Figure 3. Already in 1995, single mothers display a large and statistically significant increase of about 200 yearly hours worked relative to the control group. Results are similar for employment, with a statistically significant six percentage point increase in the probability of being employment for exposed single mothers relative to single women, representing an eight percent change in employment probability relative to the pre-reform mean. The pre-trends have not qualitatively changed relative to our baseline analysis in Figures 4-a and 4-b.[26]

Altogether, Figure 4 reveals that the finding of an insignificant effect of the EITC on labor supply depends on the inclusion of a large number of untreated individuals in the treatment group. Despite the strong positive marginal effects of the EITC on labor supply we discussed before, the zero empirical effect estimated via DiD is driven by the fact that many single mothers do not qualify for the EITC.

4 A Static Model of Labor Supply with Social Programs

The previous section conducted *ex post* analysis of the welfare-to-workfare transition as it occurred in the mid-1990s. We now use a static structural labor supply model to conduct *ex ante* counterfactual analysis of similar transitions in new environments characterized by different labor market incentives. As one of our goals is to evaluate reforms to the EITC and welfare under different tax and transfer regimes, we will explicitly model the tax and transfer system as well as the EITC, welfare, and food stamps (SNAP) in the agent's budget constraint.

[26]The analysis displays similar results if single mothers unaffected by the policy reform, e.g. earnings above the EITC threshold, were assigned to the control group. As a further test, Appendix Figure C-1 replicates the analysis on the subsample of single mothers and single women without children reporting labor income below the EITC eligibility threshold of $30,000. The analysis displays the absence of differential trends through 1993 by treatment status and sizable positive treatment effects in the post-reform period for both hours worked (Figure C-1-a) and employment status (Figure C-1-b).

In this model, single mothers have preferences over consumption, $c_{i,t}$, and hours worked, $h_{i,t}$, given by

$$u_i(c_{i,t}, h_{i,t}) = \frac{1}{1-\frac{1}{\eta}} c_{i,t}^{1-\frac{1}{\eta}} - \frac{\alpha_i}{1+\frac{1}{\gamma}} h_{i,t}^{1+\frac{1}{\gamma}} \ . \tag{10}$$

The parameter $\eta > 0$ captures the curvature of utility with respect to consumption, where a higher value indicates less concavity, while $\gamma > 0$ is the Frisch elasticity of labor supply. It determines the elasticity of hours worked with respect to changes in wages, holding the marginal utility of consumption fixed. In each period t, single mothers decide how much to work by solving

$$\begin{aligned} &\max_{c_{i,t}, h_{i,t}} \ u_i(c_{i,t}, h_{i,t}) \\ \text{subject to } c_{i,t} = &\omega_{i,t} \cdot h_{i,t} - Tax_{i,t}(\omega_{i,t} \cdot h_{i,t}) + EITC_{i,t}(\omega_{i,t} \cdot h_{i,t}) + \\ &SNAP_{i,t}(\omega_{i,t} \cdot h_{i,t}) + T_{i,t}^{Welfare}(\omega_{i,t} \cdot h_{i,t}) \ , \quad h_{i,t} \geq 0 \ , \quad c_{i,t} \geq 0 \ . \end{aligned}$$

We allow for heterogeneity in the disutility of working, $\alpha_i > 0$, so that the model can generate self-selection into employment and hours worked on the basis of unobservables. We characterize the various tax and transfer programs as part of the budget constraint. First, we model the three main existing programs: (i) the Earned Income Tax Credit (EITC); (ii) Aid to Families with Dependent Children/Temporary Assistance for Needy Families (AFDC/TANF); and (iii) the Supplemental Nutritional Assistance Program (SNAP). We use the benefit formulas for each of the three programs to recover the correct level of benefits claimed by each household. Second, we model the tax on labor income $Tax_{i,t}(\cdot)$ via a parametric function that maps pre-tax labor income to after-tax labor income. This approach is relatively common in the public finance and labor literature (see for example Benabou 2002; Guner, Kaygusuz, and Ventura 2014; Blundell, Pistaferri, and Saporta-Eksten 2016; Heathcote, Storesletten, and Violante 2017; Holter, Krueger, and Stepanchuk 2019). It allows us to independently parameterize the level and the progressivity of the tax system as follows:

$$\omega_{i,t} \cdot h_{i,t} - Tax_{i,t}(\omega_{i,t} \cdot h_{i,t}) = \theta_{0,s,t,k} \cdot (\omega_{i,t} \cdot h_{i,t})^{1-\theta_{1,s,t,k}} \ , \tag{11}$$

where we allow the tax function to vary by state s, year t, and number of children k. The parameters $\theta_{0,s,t,k} \geq 0$ and $\theta_{1,s,t,k} \in [0,1]$ capture the take-home rate and the progressivity of the tax on labor income, respectively. A higher value of $\theta_{0,s,t,k}$ implies a higher take-home rate (lower level of tax rates), while a higher value of $\theta_{1,s,t,k}$ implies greater progressivity.

In every period t, each individual receives a wage offer $\omega_{i,t}$. The initial wage offer (ω_0) is distributed according to a conditional log-normal distribution,

$$\ln \omega_0 \sim N(\mu_\omega, \sigma_\omega | k) \ , \tag{12}$$

which we allow to vary by the number of children. We also model the evolution of the log-wage offer as a random walk:

$$\ln \omega_{i,t} = \ln \omega_{i,t-1} + \nu_{i,t} \quad \text{with} \ \ \nu_{i,t} \sim N(0, \sigma_\nu^2) \ , \tag{13}$$

where the innovation, ν_t, is assumed to be normally distributed among individuals with mean zero and standard deviation σ_ν.

5 Estimation

We estimate the model via simulated method of moments (SMM). Denote the set of moments we are trying to match by M and the set of model parameters by $\Omega = \{\eta, \gamma, \{\alpha_i\}_i, \mu_\omega, \sigma_\omega^2, \sigma_\nu^2\}$. Given a wage offer and a particular tax and transfer regime, we simulate each individual's optimal labor supply choice. We then use the data created by these simulated choices to construct a set of moments, M_S, analogous to the moments, M, observed in the data. We estimate the model using the ASEC data from the pre-reform period (through 1993), while the 1995-1996 data from the transition period is used to validate the model. Our SMM estimator is

$$\widehat{\Omega} = \arg\min_{\Omega} \left(M - M_S(\Omega)\right)' W \left(M - M_S(\Omega)\right) \ , \tag{14}$$

where W is a positive semidefinite weighting matrix.[27] In practice, we set the weighting matrix equal to the inverse of the covariance matrix of the moments, $W = \Sigma_M^{-1}$, with Σ_M determined by 100 bootstrap replications of the data set. We target the following 17 moments in the data from 1988-1993 to recover 10 parameters: mean hours and employment at the aggregate level and by number of children, the mean and standard deviation of accepted wages, the autocovariance of log wages at the aggregate level and by number of children, and the causal *ex post* effects of EITC and welfare on hours worked.[28]

We model the disutility parameter according to the equation

$$\alpha_i = \alpha_k + \alpha \cdot \upsilon_i \ , \tag{15}$$

which allows the disutility of labor to vary with the number of children, k. $\upsilon_i \sim unif\, \{0.1, 2.5\}$ is a discrete uniform random variable taking six equally-spaced values between 0.1 and 2.5. This parsimonious approach fits the data well by allowing α_i to take on 18 possible values with only four underlying parameters.

5.1 Parameter Estimates

Preferences. Table 3 shows the estimates for the preference parameters. We estimate a relatively high curvature for the utility over consumption ($\eta = 0.67$) and a Frisch elasticity of one ($\gamma = 1.03$), which is consistent with the previous findings in the literature of a high elasticity of labor supply among single mothers (Attanasio et al. 2018, Blundell et al. 2016, Keane and Rogerson 2012, and Keane and Rogerson 2015).[29]

The disutility of hours worked is heterogeneous around a point estimate of $\alpha = 0.072$. This translates to a mean value for mothers with two children of 0.09, with

[27] Two reasons drive the choice of the simulated method of moments with respect to a likelihood-based method. First, the SMM approach overcomes the additional source of computational burden which arises from the multi-dimension integration problem associated with the maximum-likelihood estimator of this model. Second, we believe this method highlights more transparently the identifying variation of our model, as it allows us to replicate the causal regression coefficients of the effect of EITC and AFDC/TANF benefits on hours worked.

[28] We target the regression coefficents from the specification in column (6) of Table 1 in a sample comprising only the years prior to the 1993 reform.

[29] The estimate of η would be equivalent to a coefficient of relative risk aversion of approximately 1.5, although we do not have any source of risk in our framework.

the range of values that goes from 0.007 to 0.181.[30] The mean disutility of hours worked is 0.48 percent lower for mothers with one child and 1.6 percent higher for mothers with three children or more.

Wage Process. Table 4 presents estimates for the preference parameters. We find that the mean log-wage offer is $\mu_{\omega,k=1} = 2.56$ for mothers with one child. The average log-wage offer decreases monotonically with the number of children. Mothers with two children receive wage offers that are on average 0.04 log-points lower. The average wage offer is 0.26 log-points lower for mothers with three children or more.

Finally, we estimate a fairly large dispersion for the the unobserved heterogeneity in the initial wage offer, with a standard deviation of $\sigma_\omega = 0.582$. To put this value in perspective, the estimated standard deviation is more than twice the difference in mean log-wage offers between mothers with one child and mothers with three children or more. Blundell, Pistaferri, and Saporta-Eksten (2018) find a standard deviation of 0.533, although the authors focus on women in intact families in the Panel Study of Income Dynamics (PSID). Moreover, we find a small role of the stochastic innovation in determining the evolution of the wage offers, with an estimated standard deviation of $\sigma_\nu = 0.018$.

5.2 In-sample Fit

Tables 5 and 6 suggest that the model is successful in replicating the targeted moments. In each table we report both the data moments, M, as well as the simulated moments, M_S, calculated at the model solution.

Table 5 reports three panels. Table 5-a shows that the model replicates the mean number of hours worked as well as the negative gradient in hours worked per child during the years 1988-1993. Table 5-b shows that the model also replicates the aggregate employment rate (about 0.76) and the negative gradient of employment with respect to the number of children. Mothers with one child are the most likely to work, with an employment rate of 0.84, but employment rates drop to

[30] Given our assumption that $\upsilon \sim unif\{0.1, 2.5\}$ and the point estimate of $\alpha = 0.072$, we get $E[\alpha_i] = \alpha \cdot \frac{0.1+2.5}{2} = 0.09$, while the minimum and maximum values are $\alpha \cdot 0.01 = 0.007$ and $\alpha \cdot 2.5 = 0.181$.

0.77 and 0.61 for mothers with two and three or more children, respectively. Table 5-c shows the in-sample fit for accepted wages, an endogenous object in the model. The model successfully replicates the first and second moments of the accepted wage distribution, although the persistence of accepted log-wages is higher in the model than the data.

Table 6 shows the model fit for the causal *ex post* regression coefficients of the effects of EITC and welfare on hours worked. The model replicates the positive effect of EITC benefits on hours worked, as well as the negative effect of welfare. During the 1988-1993 period, a \$1,000 increase in EITC benefits causes an average increase of about 175 hours worked per year, while the same increase in welfare benefits causes an average reduction of 12 hours per year. Although the model slightly overstates the marginal effect of EITC on hours, the difference is not statistically significant.

5.3 Out-of-sample Predictions

Before examining the model's out-of-sample fit, we must take a stand on the wage offer equation in 1995-1996, a task complicated by the fact that we only observe the distribution of accepted wages in the data. We deal with this challenge by adding two features to the model: a new set of parameters for the wage offer model characterized by (12) and (13) and a utility cost of working, χ_i. The utility cost of working modifies preferences as follows:

$$u_i(c_{i,t}, h_{i,t}) = \frac{1}{1-\frac{1}{\eta}} c_{i,t}^{1-\frac{1}{\eta}} - \frac{\alpha_i}{1+\frac{1}{\gamma}} h_{i,t}^{1+\frac{1}{\gamma}} + \chi_i \mathbb{1}(h_{i,t} > 0) \ .$$

χ_i follows the same distribution of the disutility of working (α_i): $\chi_i = \chi \cdot \upsilon_i$, where χ is a free scale parameter. This specification allows for a "cohort-specific" cost of working and can capture heterogeneity in the unobserved cost of working that is unrelated to hours and not reflected in wages.[31]

[31] Because the extensive margin decision is an endogenous choice in the model, we want to avoid estimating the wage offer equation outside the model via some parametric reduced-form models that could be inconsistent with our structural model. An alternative estimation strategy would pool together data from the pre-reform and transition periods and estimate the model by allowing the wage offer parameters and the cost of working to vary by time period. We prefer our estimation strategy as it allows us to use moments from the 1995-1996 data to test the model.

We calibrate χ and the new wage offer model by only matching moments of the accepted wage distribution in 1995-1996. The rest of the preference parameters remain at their estimated values presented in Table 3. This approach lets us test if the preferences we estimate using data prior to 1994 can replicate labor supply statistics that we do not directly target in 1995-1996.

Figure 5 compares the average hours worked in the data and in the model. In each graph, the first set of bars show the in-sample fit already discussed in Table 5-a. The second set of bars shows the performance of the model for the untargeted out-of-sample moments. Overall, the model predicts an increase in hours worked during the period of interest that is consistent with the data. The rise in hours worked appears both at the aggregate level (Figure 5-a) and by number of children (Figures 5-b, 5-c, and 5-d). The validation exercise for employment rates in Figure 6 suggests similar results.

6 *Ex Ante* Counterfactual Analysis

We use the estimated model to analyze three counterfactual reforms. First, we examine the effect of reforms to each program (either the EITC or welfare) during the welfare-to-workfare transition in isolation, so that we can disentangle each program's effect on labor supply. Then we consider two new policies: a large expansion of the EITC in 1996 and the replacement of either the EITC or welfare with Universal Basic Income. We show that the response of labor supply to these two proposed policies varies considerably with the progressivity of the tax code.

6.1 EITC and Welfare Reform Decomposition

Our first counterfactual exercise consists of several scenarios. The first keeps all tax and transfer programs at their 1996 level with the exception of the EITC, whose benefits are artificially held at their level in 1993. The second scenario repeats the exercise for welfare by combining AFDC's policies from 1993 with the 1996 tax and transfer system. The third scenario keeps both EITC and welfare at their 1993 policy rules, while leaving the rest of tax system at its 1996 level.

Figure 7 presents the results of this exercise. We find that in the absence of reforms to either the EITC or welfare, both employment and hours worked would have been lower. Employment for single mothers would have been 7.5 percentage points lower if the EITC were not expanded as it was, or 5.5 percentage points lower if AFDC were not reformed as it was. Hours worked would have dropped by 20, respectively 48, hours per year if either the EITC were not expanded or welfare not reformed as they were through 1996.

The two reforms combined to reduce the 1996 employment rate by 12 percentage points and mean yearly hours worked by 73 hours for single mothers. This is an interesting result as the model suggests that the counterfactual level of employment and hours would have decreased after 1993 due to rises in SNAP benefits and changes in wage offers and preferences for working. This finding is consistent with the fact that, through 1993, both employment and hours were trending downward for single mothers, and the model suggests that these would have kept falling without the reforms.

6.2 The Interaction of Social Programs and the Tax Code

Our remaining counterfactual exercises demonstrate how the exact same reform can generate different effects on labor supply depending on the tax regime and method of financing the reform. Denote a specific social program by a vector of parameters, Υ_j, that fully characterize the program, and denote the entire tax and transfer system by a set of social programs: $\Upsilon = \{\Upsilon_1, \ldots, \Upsilon_J\}$. We show that the labor supply response to a single program, Υ_j, can depend on the other taxes and transfers already in place, Υ_{-j}.

We evaluate a reform to program j $\left(\Upsilon_j' \neq \Upsilon_j\right)$ using a standard potential outcomes framework, where for each individual i we define the pair of potential outcomes, $\left(\mathrm{Y}_i(\Upsilon_j', \Upsilon_{-j}), \mathrm{Y}_i(\Upsilon_j, \Upsilon_{-j})\right)$. We then examine the reform's average treatment effect (ATE):

$$ATE_j(\Upsilon_{-j}) = E\left[\mathrm{Y}_i(\Upsilon_j', \Upsilon_{-j}) - \mathrm{Y}_i(\Upsilon_j, \Upsilon_{-j})\right] . \qquad (16)$$

We use our estimated model to simulate changes in the policy regime Υ_j and observe the entire distribution of potential outcomes, $\left\{\mathrm{Y}_i(\Upsilon_j', \Upsilon_{-j}), \mathrm{Y}_i(\Upsilon_j, \Upsilon_{-j})\right\}_i$, for a fixed parameterization of the tax code ($\theta_{0,s,t,k}$ and $\theta_{1,s,t,k}$ in Equation (11)).

We first examine the response of the extensive margin to an expansion of the EITC that increases maximum benefits by $2,000 and the federal income limit by $5,000 in 1996, while simultaneously varying the tax on labor income. Figure 8-a shows how the ATE of the reform depends on the take-home rate, θ_0. The *x-axis* is defined relative to the original 1996 take-home rate, which means that a level of 1.1 is a take-home rate that is ten percent higher than baseline.[32] ATEs are higher when take-home rates are lower (tax rates are higher). For example, the ATE on employment goes from 7 to 2.5 percentage points if we vary the take-home rate from 0.8 to 1.2 times the baseline 1996 level. This result is sensible as the marginal benefit of a tax credit like the EITC depends on the tax rate, with workers facing an extensive margin decision benefitting more from the credits when income taxes are higher.

Figure 8-b shows a similar pattern for the progressivity of the tax system, detemined by θ_1: The ATE is higher when individuals face steeper marginal tax rates as their income rises (θ_1 increases). These results suggest that the estimated ATE of the EITC on employment is not invariant to changes in taxes. However, a large body of the empirical public economics literature (e.g. Saez 2002 and Eissa, Kleven, and Kreiner 2008) uses *ex post* evaluation methods to recover the average extensive margin elasticity for the population of interest. Figure 9 demonstrates that this elasticity varies widely according to the tax system. Each dot represents the average percentage change in the probability of being employed induced by a one percent change in the take-home rate. The wide variation in estimated elasticities is caused by changes in the composition of individuals at the margin of employment as the tax code changes.[33]

Our final counterfactual exercise examines the effects of Universal Basic Income on the labor supply of single mothers. UBI has recently gained traction in policy circles as a result of the dislocations created by rapid technological development

[32] Tax rates are heterogeneous by state and number of children. We proportionally change the various tax rates by the same factor (x-axis).

[33] Similarly, Attanasio et al. (2018) find that the aggregate elasticity is not a structural parameter, as it varies over the the business cycle because of the heterogeneity of marginal individuals in different aggregate states of the economy. Moreover, Moffitt (2019) finds that the marginal treatment effects of welfare reforms on labor supply change over time because of the preference heterogeneity of the marginal individual.

and COVID-19. We replace either the EITC or welfare with a UBI program that targets the population of single mothers in a budget-neutral way.[34] Table 7 shows the effect of UBI on employment and hours worked. Contrary to common wisdom, UBI can generate positive or negative effects on labor supply depending on the program it replaces. If UBI completely replaces the EITC, employment and hours worked would fall by 24 percentage points and 226 hours per year, respectively. While these results are large, the considered reform is also massive, pooling all of the EITC money received by single mothers in 1996 and redistributing it in (approximately) $3000 checks to each individual. If UBI instead replaces welfare, the effect on hours and employment is positive. Given our *ex post* analysis, this finding is hardly surprising: The reform removes a disincentive to enter the labor force by eliminating benefits targeting people who do not work and equalizes the cash transfer to everybody regardless of employment status.

Figure 10 shows that the responses of hours worked to the UBI reform are heterogeneous and non-monotone. When UBI replaces welfare, shown in Figure 10-a, there are strong positive effects on labor supply for individuals at the lowest quintile of the income distribution, those who directly lose access to welfare benefits. The rest of the population reduces labor supply as a consequence of the unconditional transfer. On the other hand, when UBI replaces the EITC, we find opposing intensive margin effects depending on whether the individual was either at the phase-in or at the phase-out of the EITC schedule. Individuals at the phase-in lose the negative marginal tax rates on earnings, causing a reduction in hours worked, while individuals at the phase-out experience the opposite. These two forces explain the non-monotone effects in Figure 10-b.

[34] Kearney and Mogstad (2019) provide a review of the UBI proposals and highlight that they risk being extremely expensive and inefficient, e.g. through increasing income inequality. We adopt a concept of UBI that resembles the one in Hoynes and Rothstein (2019), namely that the cash transfer is universally provided to the whole population of single mothers but, due to the budget neutral criterion to finance it, it is not sufficiently generous to allow recipients to live on it without additional earnings. On the latter aspect, our definition is similar to the one in Banerjee, Niehaus, and Suri (2019). Jones and Marinescu (2020) study the long-run impact of the Alaska Permanent Fund Dividend, a universal and unconditional cash transfer program introduced in 1982 and find that the program did not affect long-run aggregate employment (extensive margin) but increased the share of Alaskans working in part-time jobs.

7 Conclusions

The goal of evidence-based policymaking is to learn from past reforms (*ex post*) to predict the effects of future policy regime changes that never happened before (*ex ante*). Any *ex post* analysis that seeks to accomplish this goal must use methods that can identify whether individuals respond to the incentives generated by the social program being studied. In this paper we show that the oft-used DiD design cannot answer this question when it is applied to reforms of pre-existing social programs with continuous levels of treatment. We propose a different approach that can identify the *marginal* responses of labor supply to social program benefit levels. Our *ex post* analysis finds significant marginal responses of the labor supply of single mothers to both the EITC and welfare reforms of the 1990s, with results mostly driven by extensive margin responses, as well as intensive margin responses of individuals at the phase-in of the EITC schedule.

We use our estimated causal *ex post* evaluation of EITC and welfare reforms to identify a model of labor supply with heterogeneous exposure to multiple tax and transfer programs. The estimated model shows that the effect of an additional EITC expansion on employment heavily depends on the level and progressivity of labor income taxes. This result suggests that the evaluation of past EITC reforms does not speak for itself about the effects of future EITC expansions. The evolution of the tax code over time affects the choices made by individuals, which in turn determine the aggregate labor supply response. For this reason, without careful modeling of the incentives faced by individuals, *ex post* analysis provides limited guidance to policymakers considering new reforms.

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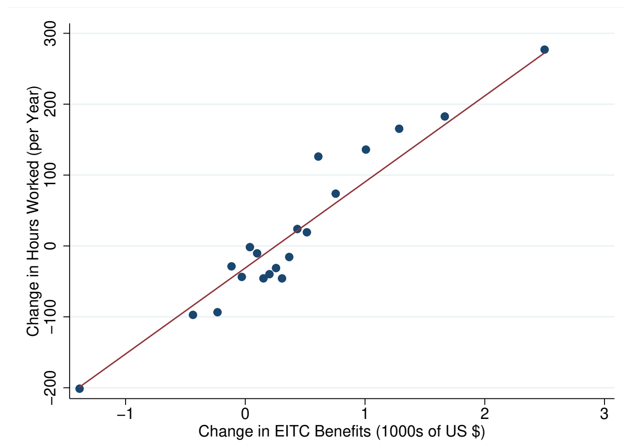
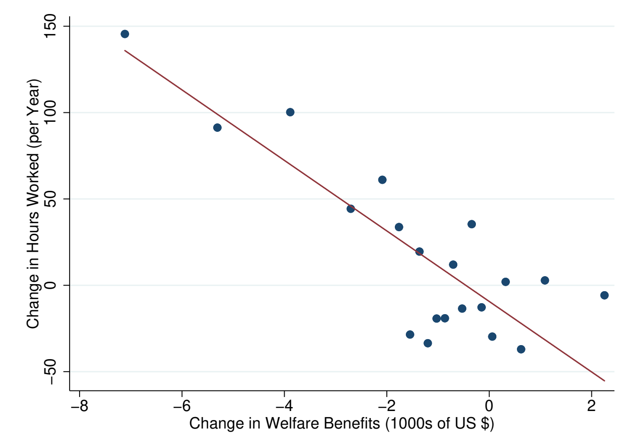
Figure 1: Hours Worked, Welfare Transfers and EITC Benefits



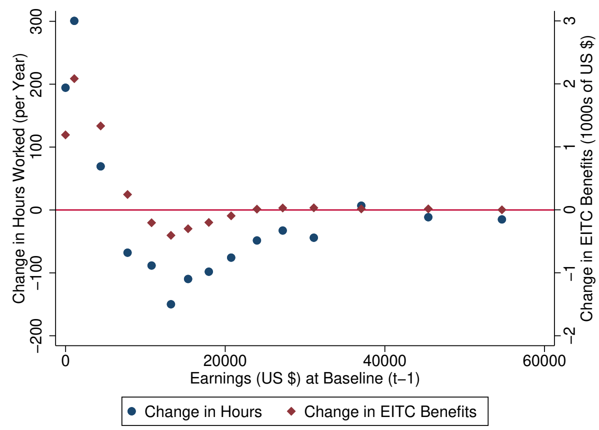
(a) Hours and EITC



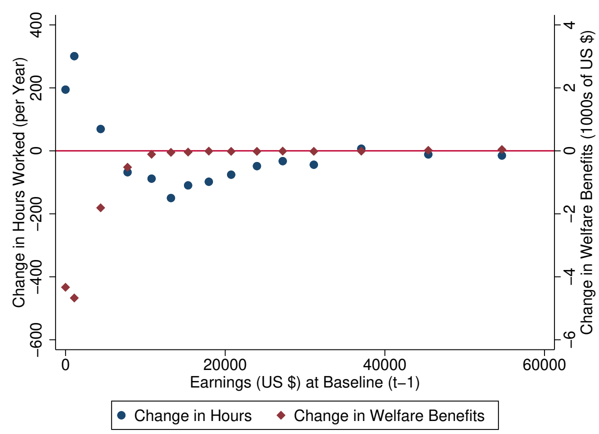
(b) Hours and Welfare

The figure shows the relation between policy-induced changes in EITC benefits (Panel (a)) and welfare benefits (Panel (b)) on the change in yearly hours worked by single mothers. Policy-induced changes in EITC and welfare benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC and welfare benefits are provided in Section 3.1 and Appendix D. Each panel depicts the binscatter of the nonparametric relation and linear fit line between the y-residuals on x-residuals with specifications containing control variables for mother's race, number of dependent children (indicators), year fixed effects, state fixed effects, state unemployment level and state welfare waivers (indicator).

Figure 2: Distribution of Changes in Hour Worked, EITC and Welfare



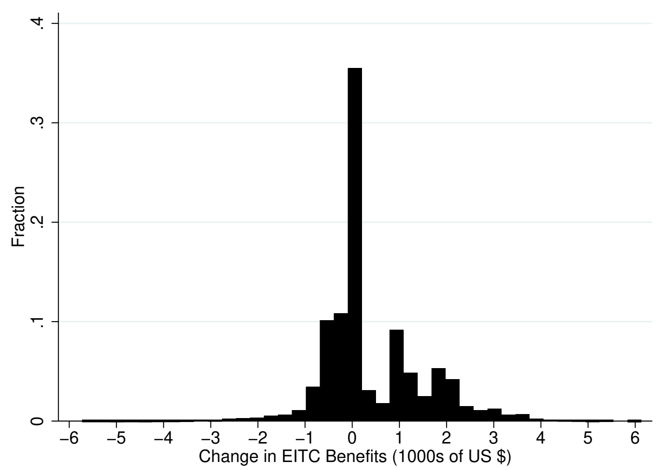
(a) Hours Worked and EITC



(b) Hours Worked and Welfare

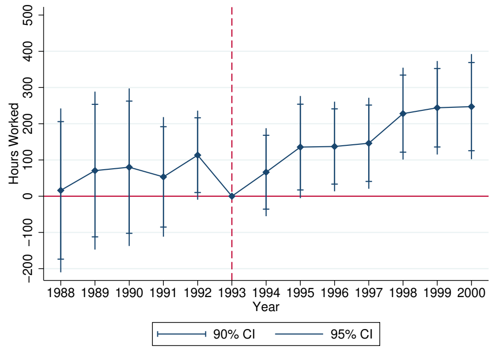
The figure shows the relation between policy-induced changes in EITC benefits or welfare benefits, change in yearly hours worked by single mothers, and earnings at baseline ($t-1$). Panel (a) illustrates the analysis of policy-induced changes in EITC benefits. Panel (b) illustrates the analysis of policy-induced changes in welfare benefits. Policy-induced changes in EITC and welfare benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC and welfare benefits are provided in Section 3.1 and Appendix D. Each panel depicts the binscatter of the nonparametric relation between the year-on-year change in hours worked by single mothers (left y-axis and blue dots) and earnings at baseline and the relation between the policy-induced change in EITC (Panel (a)) or welfare (Panel (b)) benefits (right y-axis and red diamonds) and earnings at baseline.

Figure 3: Distribution of EITC Benefit Changes for Single Mothers in the ASEC

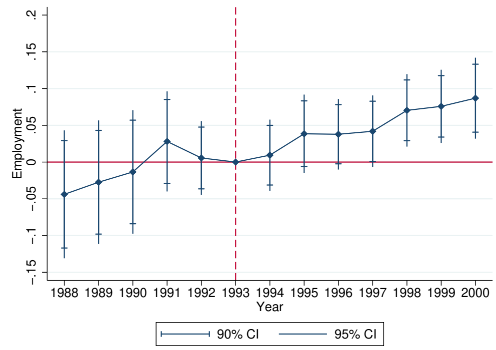


The figure shows the distribution of policy-induced changes in EITC benefits for single mothers in the CPS-ASEC. Changes in EITC benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC benefits are provided in Section 3.1 and Appendix D.

Figure 4: Event Study Analysis of the 1993 EITC Reform

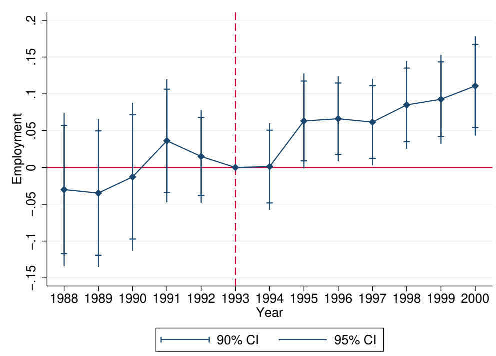


(a) Hours: Baseline Analysis



(b) Employment: Baseline Analysis

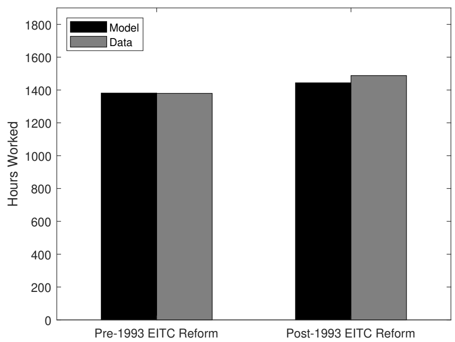
(c) Hours: "As Treated" Analysis



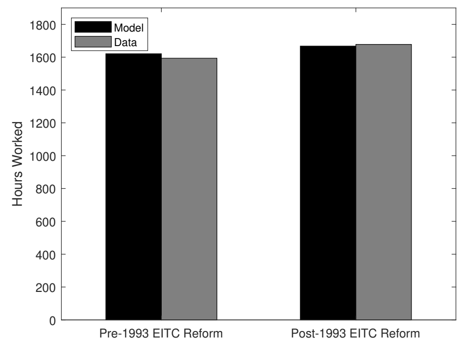
(d) Employment: "As Treated" Analysis

The figure compares two event study analyses of the effect of the 1993 EITC reform on yearly hours worked (Panels (a) and (c)) and employment (Panels (b) and (d)). In Panels (a) and (b), the treatment group is comprised of single mothers and the control group is comprised of single women without children. The specification in Panels (c) and (d) exclude single mothers who were unexposed to changes in EITC benefits caused by the 1993 EITC reform. Details on the definition of the group of single mothers unexposed to policy-induced changes in EITC benefits are provided in Section 3.3. Yearly hours worked (in Panels (a) and (c)) and an indicator variable for employment status (in Panels (b) and (d)) are regressed on a set of interaction terms of the indicator variable for the treatment group and indicator variables for each year in the period 1988-2000. The event study specification in each panel also contains control variables for the number of dependent children (indicators), state fixed effects, state unemployment level and state welfare waivers (indicator). The year of the reform's passage, 1993, is the reference year for the analysis. The red vertical line separates the pre-reform (1993 and earlier) period from the post-reform period. Each panel shows the point estimates for the treatment effect of the reform together with the 90 and 95 percent confidence intervals based on standard errors clustered at the individual level.

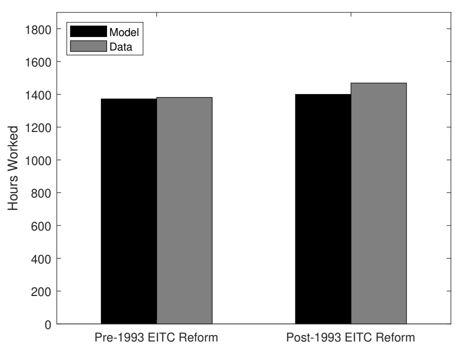
Figure 5: Validation: Predicted Hours Worked Pre- and Post-1993 EITC Reform



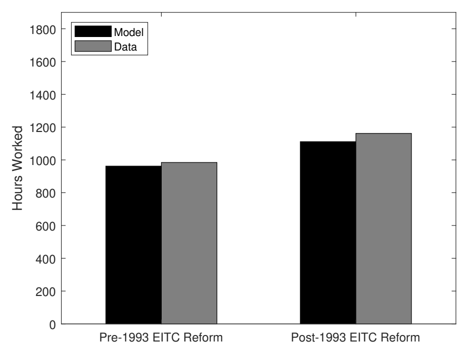
(a) Whole Sample



(b) Mothers with one Child



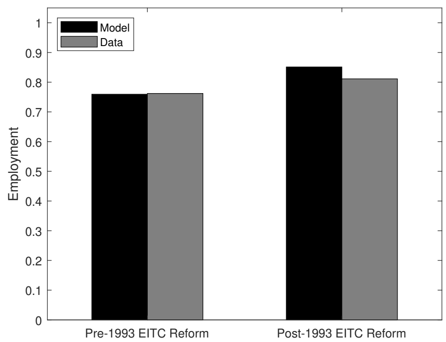
(c) Mothers with two Children



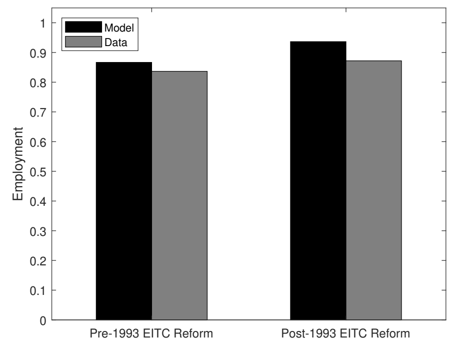
(d) Mothers with three Children

The figure shows the model's predictions for hours worked. In each graph, the first set of bars shows the fit of the model for the years 1988-1993, prior to the implementation of the 1993 EITC reform, while the second set of bars shows the performance of the model for the untargeted moments of hours worked in 1995-1996. The analysis is performed on the whole sample of single mothers (Panel (a)) and the sample of single mothers with one child (Panel (b)), two children (Panel (c)), and three or more children (Panel (d)). The figure displays yearly hours worked by single mothers as predicted by the model (black bars) and as observed in the data (gray bars).

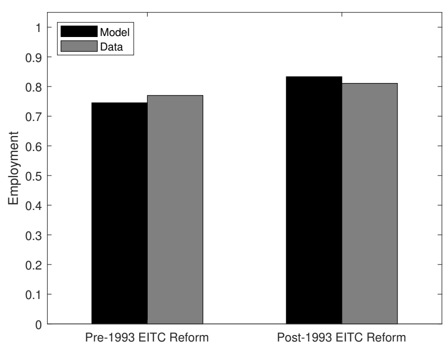
Figure 6: Validation: Predicted Employment Pre- and Post-1993 EITC Reform



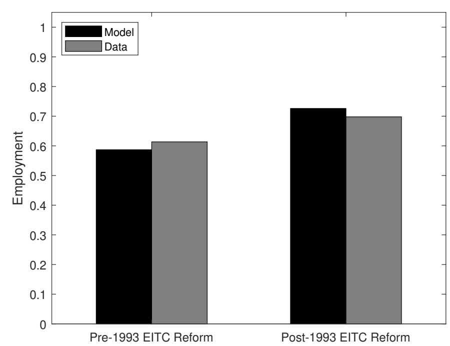
(a) Whole Sample



(b) Mothers with one Child



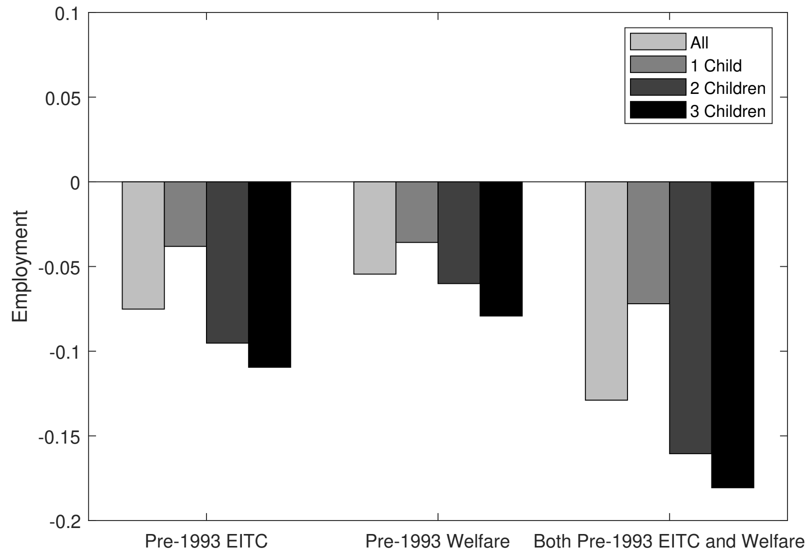
(c) Mothers with two Children



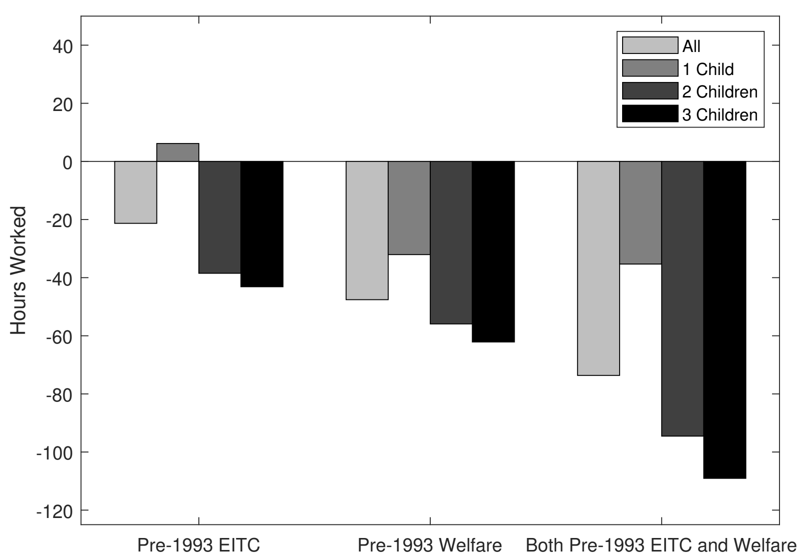
(d) Mothers with three Children

The figure shows the model's predictions for employment. In each graph, the first set of bars shows the fit of the model for the years 1988-1993, prior to the implementation of the 1993 EITC reform, while the second set of bars shows the performance of the model for the untargeted moments of employment in 1995-1996. The analysis is performed on the whole sample of single mothers (Panel (a)) and the sample of single mothers with one child (Panel (b)), two children (Panel (c)), and three or more children (Panel (d)). The figure displays the employment rate of single mothers as predicted by the model (black bars) and as observed in the data (gray bars).

Figure 7: Policy Contributions in 1996 of Pre-1993 EITC and Welfare Policies



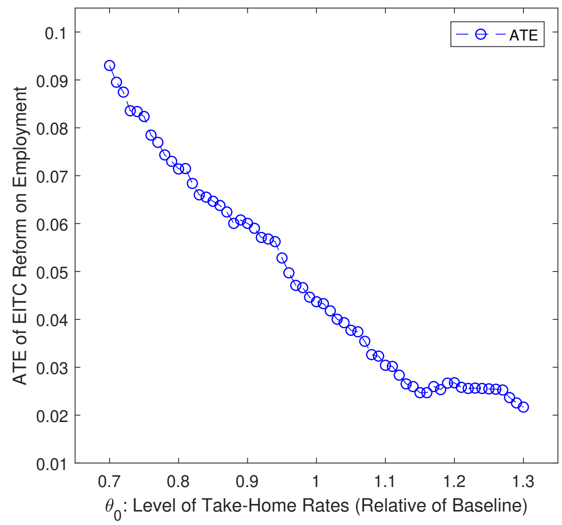
(a) Employment



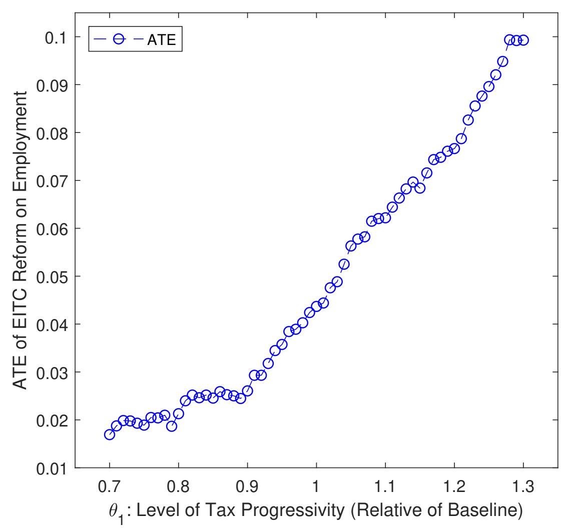
(b) Hours Worked

The figure shows the counterfactual level of hours worked (Panel (a)) and employment (Panel (b)) for single mothers in 1996 if counterfactually either the EITC or welfare were held at the 1993 regime, while keeping the rest of the other tax and transfer programs at their 1996 levels. The analysis is performed on the whole sample of single mothers, as well as by number of children.

Figure 8: Estimated ATE by Different Tax Regimes



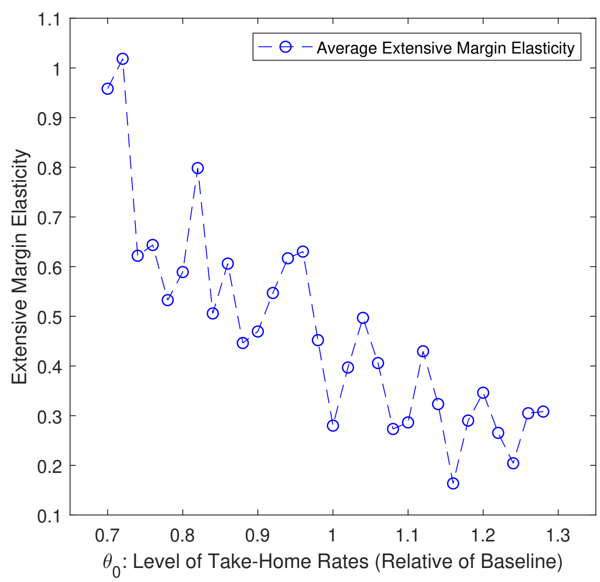
(a) ATE and Tax Rates



(b) ATE and Tax Progressivity

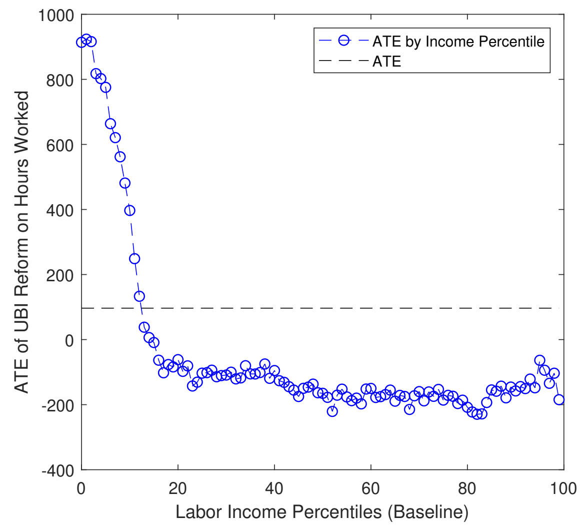
The figure shows the ATE (blue dots) of an EITC reform on employment as a function of the level (Panel (a)) and the progressivity (Panel (b)) of the tax regime. The simulated EITC reform includes an expansion of the federal exhaustion point of the EITC schedule of $5,000, as well as an increase of the maximum federal credits of $2,000.

Figure 9: Estimated Extensive Margin Elasticity by Different Tax Regimes

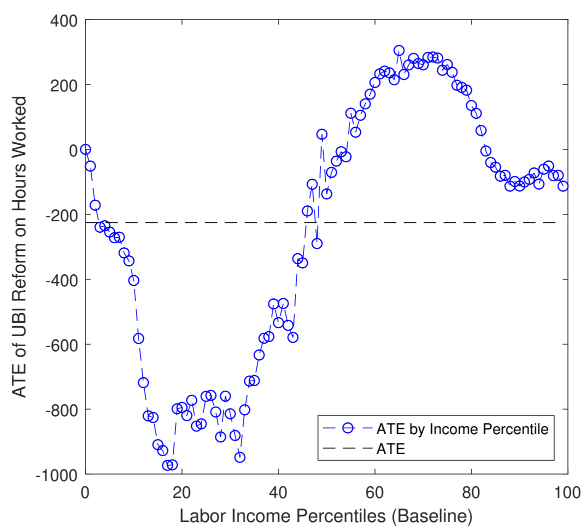


The figure shows how the estimated aggregate extensive margin elasticity to taxes varies by the level of the tax rates. Each elasticity (blue dot) is the percentage change in the aggregate employment rates caused by a small ($\frac{0.01}{\theta_0}$) change in the tax rates.

Figure 10: Heterogeneous Effects of UBI on Hours Worked by Financing Options



(a) UBI replaces TANF



(b) UBI replaces EITC

The figure shows the response of hours worked in the population to a welfare reform that substitutes UBI for TANF (Panel (a)) or EITC (Panel (b)). Each dot represents the ATE by labor income percentile at baseline. The dashed line represents the ATE in the population.

Table 1: EITC Benefits,Welfare Benefits and Hours Worked per Year

	(1)	(2)	(3)	(4)	(5)	(6)
	Outcome: Yearly Hours Worked					
Change in EITC Benefits ($ 1000s)	125.45***		109.96***	115.38***	115.04***	109.90***
	(7.28)		(8.56)	(8.89)	(8.91)	(9.51)
Change in Welfare Benefits ($ 1000s)		-31.62***	-12.08***	-17.01***	-17.16***	-16.86***
		(2.15)	(2.57)	(2.85)	(2.87)	(3.08)
N	10959	10959	10959	10959	10959	10959
Controls and State F.E.	No	No	No	Yes	Yes	Yes
Unemployment and Waiver Controls	No	No	No	No	Yes	Yes
Control Function	No	No	No	No	No	Yes

The table shows the causal effect of changes in policy-induced EITC benefits and welfare benefits on the change in yearly hours worked by single mothers. The dependent variable is the year-on-year change in hours worked by single mothers. Policy-induced changes in EITC and welfare benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC and welfare benefits are provided in Section 3.1 and Appendix D. Control variables include mother's race, indicator variables for the number of dependent children, and year fixed effects. Unemployment and waiver controls include controls for state unemployment level and indicator variables for state welfare waivers. The control function includes a set of controls for lagged labor income, lagged business income, lagged farm income, and lagged non-labor income. See Appendix D for details on the control function. Standard errors are robust to heteroskedasticity and reported in parentheses. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

Table 2: EITC Benefits,Welfare Benefits and Hours Worked per Year (by Previous Employment Status)

	(1)	(2)	(3)	(4)	(5)	(6)
	Outcome: Yearly Hours Worked					
Change in EITC Benefits ($ 1000s)	72.49***		61.27***	69.62***	69.19***	71.78***
	(21.66)		(22.16)	(23.91)	(23.98)	(24.01)
Change in Welfare Benefits ($ 1000s)		13.97***	12.29***	6.20*	6.00	5.49
		(3.40)	(3.51)	(3.75)	(3.77)	(3.93)
Working in t-1 × Change in EITC Benefits ($ 1000s)	34.95		31.43	28.95	29.16	28.54
	(23.43)		(24.49)	(25.79)	(25.83)	(25.94)
Working in t-1 × Change in Welfare Benefits ($ 1000s)		-53.12***	-28.43***	-25.52***	-25.38***	-26.44***
		(5.71)	(6.37)	(6.48)	(6.48)	(6.64)
N	10959	10959	10959	10959	10959	10959
Controls and State F.E.	No	No	No	Yes	Yes	Yes
Unemployment and Waiver Controls	No	No	No	No	Yes	Yes
Control Function	No	No	No	No	No	Yes

The table shows the causal effect by lagged employment status of changes in policy-induced EITC benefits and welfare benefits on the change in yearly hours worked by single mothers. The dependent variable is the year-on-year change in hours worked by single mothers. Policy-induced changes in EITC and welfare benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC and welfare benefits are provided in Section 3.1 and Appendix D. The variable *Working in* $t-1$ is an indicator variable taking the value of one if the mother was working at baseline, and zero otherwise. Control variables include mother's race, indicator variables for the number of dependent children, and year fixed effects. Unemployment and waiver controls include controls for state unemployment level and indicator variables for state welfare waivers. The control function includes a set of controls for lagged labor income, lagged business income, lagged farm income, and lagged non-labor income. See Appendix D for details on the control function. Standard errors are robust to heteroskedasticity and reported in parentheses. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

Table 3: Estimates for Preferences Parameters

	Preferences
Curvature of Consumption (η)	0.6716 [0.6703,0.6738]
Frisch elasticity (γ)	1.0319 [1.0236,1.0372]
Disutility of Hours Worked (α)	0.0722 [0.0705,0.0774]
Additional Disutility of Hours Worked with One Child ($\%\alpha$)	-0.0048 [-0.0066,-0.0036]
Additional Disutility of Hours Worked with Three Children ($\%\alpha$)	0.0160 [-0.0010,0.0169]

The table shows the estimated preferences parameters; see Equations (10) and (15). The 95 percent confidence intervals in brackets are calculated via 100 bootstrap repetitions. The point estimates are the averages among the bootstrap repetitions.

Table 4: Estimates for Wage Process

	Wage Process
Mean (One Child, $\mu_{\omega,1}$)	2.5591 [2.5412,2.5874]
Additional Mean with Two Children ($\mu_{\omega,k=2}$)	-0.0381 [-0.0744,-0.0333]
Additional Mean with Three Children ($\mu_{\omega,k=3}$)	-0.2601 [-0.3214,-0.2311]
Initial Standard Deviation (σ_{ω})	0.5815 [0.5659,0.5917]
Standard Deviation of Innovation (σ_{ν})	0.0183 [0.0123,0.0392]

The table shows the estimated wage process parameters; see Equations (12) and (13). The 95 percent confidence intervals in brackets are calculated via 100 bootstrap repetitions. The point estimates are the averages among the bootstrap repetitions.

Table 5: In-sample Fit for Hours, Employment and Wages

	(1) Model	(2) Data
	Panel A: Hours Worked	
Mean Hours Worked	1381.2	1379.6
Mean Hours Worked (One Child)	1621.0	1593.7
Mean Hours Worked (Two Children)	1371.9	1381.0
Mean Hours Worked (Three Children)	962.3	984.4
	Panel B: Employment Rate	
Employment Rate	0.759	0.762
Employment Rate (One Child)	0.867	0.837
Employment Rate (Two Children)	0.745	0.770
Employment Rate (Three Children)	0.587	0.614
	Panel C: Accepted Wages	
Mean Accepted Wage	16.09	16.11
SD Accepted Wage	9.62	9.60
Mean Accepted Wage (One Child)	16.46	17.02
Mean Accepted Wage (Two Children)	16.79	16.33
Mean Accepted Wage (Three Children)	13.77	13.41
Autocovariance Accepted Log-Wages	0.315	0.222
SD Accepted Log-Wage	0.56	0.64

The table shows the in-sample fit for hours worked by single mothers (Panel (a)), employment rate (Panel (b)), and accepted wage (Panel (c)). The table displays outcomes as predicted by the model (column 1) and as observed in the data (column 2). All monetary values are expressed in 2015 US dollars.

Table 6: In-sample Fit for Regression of Hours on EITC and Welfare

	(1)	(2)
	Outcome: Yearly Hours Worked	
	Model	Data
Change in EITC Benefits ($ 1000's)	214.02	175.29
Change in Welfare Benefits ($ 1000's)	-19.44	-11.74

The table shows the in-sample fit for the regression of year-on-year changes in hours worked by single mothers on policy-induced changes in EITC and welfare benefits. Policy-induced changes in EITC and welfare benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC and welfare benefits are provided in Section 3.1 and Appendix D. The table displays outcomes as predicted by the model (column 1) and as observed in the data (column 2).

Table 7: *Ex ante* Estimated Effects of UBI by Different Ways of Financing the Program

	(1)	(2)
	Different Ways of Financing UBI:	
	UBI replaces AFDC/TANF	UBI replaces EITC
ATE on Employment	0.13	-0.24
ATE on Hours	96.65	-226.03

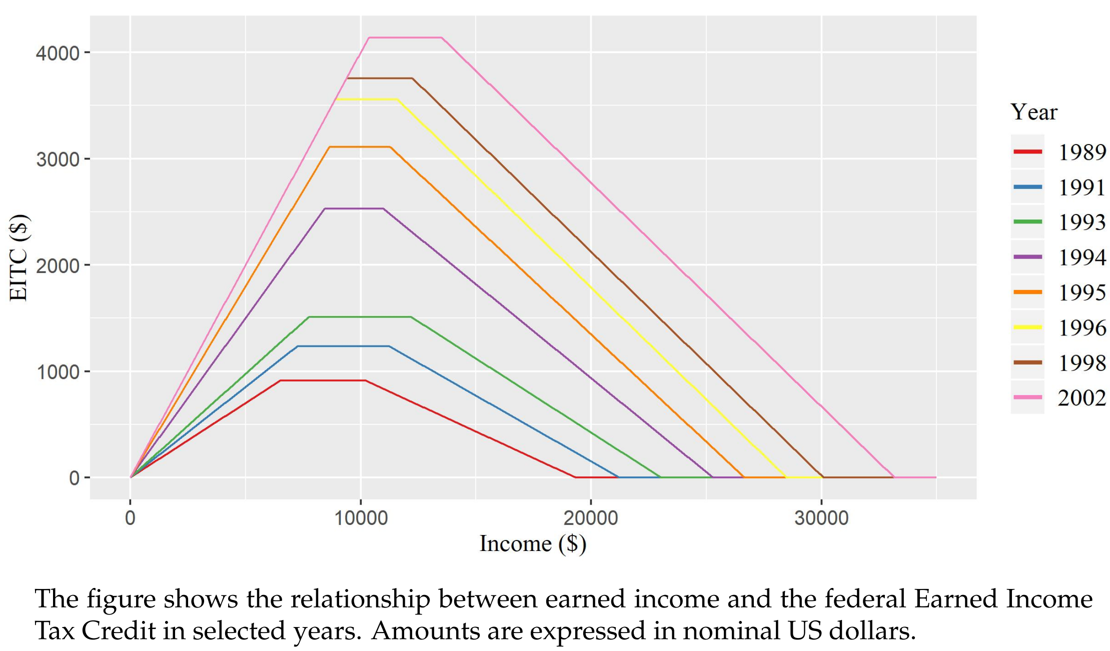
The table shows the average response of employment and hours worked in the population (ATE) to a reform that substitutes UBI for welfare (Column 1) or EITC (Column 2).

Appendices

A Institutional Background: The Welfare-to-Workfare Transition

The Earned Income Tax Credit (EITC) is one of the largest income support programs in the United States. Enacted in 1975 to provide a modest supplement to the income of working families, it has been expanded significantly by the federal government in several rounds: most notably in 1986, 1993, and 2009. While the EITC expanded in each year during the period we study, 1988-2002, Figure A-1 shows that the largest year-to-year expansions of the program occurred in precisely those years, 1994-1996, that we define to be the welfare-to-workfare transition. In the 2000s, many states also implemented and expanded their own EITC programs. Throughout the paper, we calculate an individual's EITC benefits to be the sum of federal and state EITC credits.

Figure A-1: EITC Schedule: Selected Years



The figure shows the relationship between earned income and the federal Earned Income Tax Credit in selected years. Amounts are expressed in nominal US dollars.

Unlike the EITC, welfare has historically provided benefits to mothers who do not work. In response to growing concerns that welfare's incentives were contributing to high unemployment and out-of-wedlock births, many states implemented reforms between 1992 and 1996. These so-called welfare waiver reforms contained a mix of punishments and incentives to get mothers off of the welfare rolls and into employment. They had five main characteristics: time limits, changes in exemptions from program requirements, sanctions for recipients who violated program requirements, family caps, and earnings disregards.

While Aid For Families with Dependent Children (AFDC) did not impose limits on how long beneficiaries could receive welfare, many of the welfare waivers restricted the receipt of benefits to specific periods of time, such as 24 out of every 48 months. There were three types of time limits. "Termination" time limits resulted in the loss of benefits after the limit had been reached, while "reduction" time limits caused a reduction in benefits, and "work requirement" limits did not cut off aid so long as the beneficiaries complied with state-stipulated work requirements. These time limits were not retroactive. As a result, very few people were kicked off the welfare rolls before 1997. Appendix E documents the earliest possible date that a welfare recipient could lose eligibility due to the expiration of time limits.

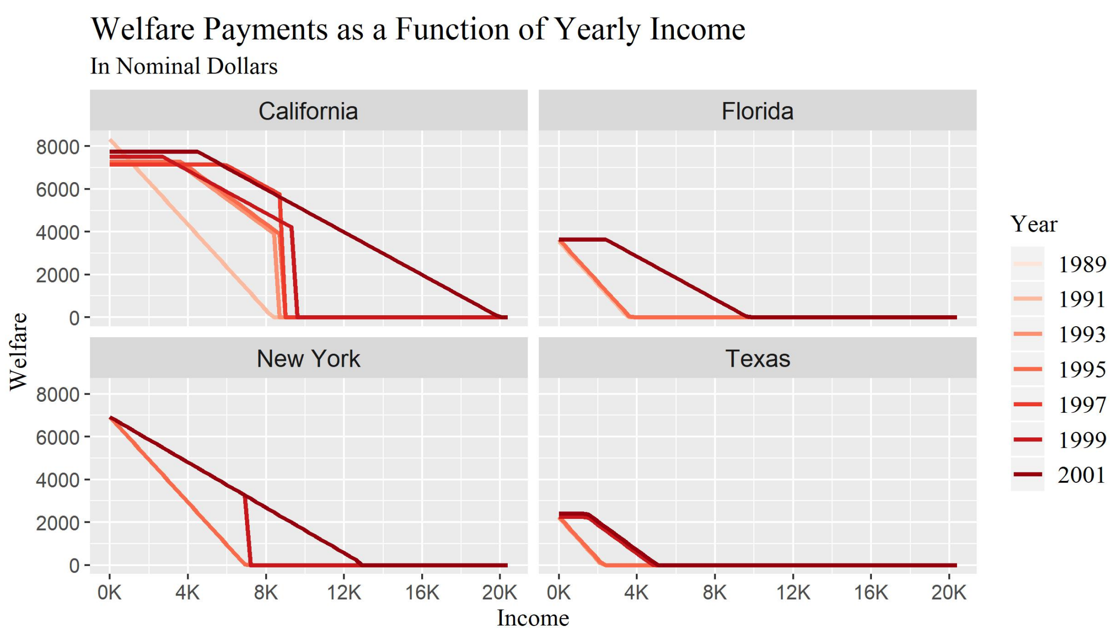
AFDC required states to run education and jobs training programs (JOBS) for welfare recipients, and partipication in JOBS (or similar activities such as secondary education and job search) was mandatory for nonexempt individuals. Federal policy exempted recipients if their youngest child was under the age of 3, but many state waivers lowered the age exemption and imposed sanctions on individuals who violated this requirement. In the most severe cases, repeated violations could result in the lifetime termination of benefits.

While AFDC stipulated that benefits increase with the number of children, several states instituted family caps that froze benefit levels if a recipient had a child while currently receiving welfare.

Lastly, under AFDC, welfare recipients faced a 100 percent marginal tax rate: Benefits were reduced one-to-one with each dollar earned through employment. Many state waivers countered this disincentive to work in two ways. Earnings *dollar* disregards allowed recipients to earn a fixed amount of money before the benefits were reduced, while earnings *rate* disregards reduced the marginal tax rate on remaining earnings to below 100 percent. Many states implemented both dollar and rate disregards simultaneously. Michigan, for example, disregarded the first $200 of monthly income and lowered the marginal tax rate to 80 percent on the remaining income. Unlike the previous four characteristics of the waivers, disregards represent the use of carrots, rather than sticks, to provide incentives for welfare recipients to participate in the labor market.

Figure A-2 displays welfare's schedule of benefits for a mother with 2 children and no nonlabor income in four large states: California, Florida, New York, and Texas. Earnings dollar disregards introduce kinks in the schedule, while earnings rate disregards reduce the magnitude of the slope.

Figure A-2: Welfare Functions by State and Year



The first states to implement the welfare waivers – New Jersey, California, and Michigan – did so in 1992. However, the majority – 23 of 30 statewide reforms – were implemented between 1994 and 1996, precisely when the EITC experienced the most dramatic expansions.

The welfare waivers culminated in the passage by Congress of the Personal Work

Opportunity and Reconciliation Act of 1996 (PRWORA). PRWORA replaced AFDC with TANF. Under TANF, each state received a block grant and was given substantial leeway in designing its TANF program. All five of the main characteristics of the statewide waivers found their way into PRWORA. States were no longer required to provide more benefits to larger families, and family caps were implemented in many states. A federal lifetime limit of five years on the receipt of benefits was instituted, although many states imposed more stringent limits. Earned income disregards became the rule rather than the exception, and many states adopted graduated sanctions, some of which could ultimately result in the lifetime loss of benefits. Many of the states that implemented waiver reforms retained these policies as part of their TANF programs. Others made modifications.

Parceling out the effects of each of the five types of welfare waiver reforms on labor supply is a difficult task and one that we do not pursue in this paper. However, it at least seems likely that the effects of the welfare waiver reforms on labor supply between 1994 and 1996 stressed in Kleven (2020) were not caused by lifetime limits. The welfare waivers imposed lifetime limits on benefits that were not retroactive, meaning that few people were kicked off the welfare rolls prior to 1997 (see Appendix E and US Department of Health and Human Services (1997)). The disregards took effect earlier and could account for increased labor supply during this period. However, in reducing the effective tax on earned income, they made welfare operate more like the EITC in its use of financial incentives designed to draw beneficiaries into the labor force.

B Proofs and Special Cases of DiD Estimand

Proof of Proposition 1.

We consider a two-period model and define an indicator variable $D_{0,0}$ that equals 1 if an individual receives no subsidy in either period and 0 otherwise. The variable denoting treatment is then given by $Treat = 1 - D_{0,0}$. We define the following terms: for $j = 1, \dots, J$, $p_{j,t} \equiv \mathbb{P}(D_{j,t} = 1)$, $p_{j,t-1} \equiv \mathbb{P}(D_{j,t-1} = 1)$, and $p_{0,0} \equiv \mathbb{P}(D_{0,0} = 1)$. Note that $1 - p_{0,0} = p_{Treat}$.

The DiD estimand is

$$\beta_3^{DiD} = \frac{cov(Y_t - Y_{t-1}, 1 - D_{0,0})}{var(1 - D_{0,0})} \ . \tag{B-1}$$

The denominator in B-1 is equal to $p_{0,0}(1 - p_{0,0})$. The numerator can be written as

$$\begin{aligned} cov(Y_t - Y_{t-1}, Treat) &= cov(Y_t - Y_{t-1}, 1 - D_{0,0}) \\ &= cov(Y_{0,t} - Y_{0,t-1}, 1 - D_{0,0}) \\ &+ cov(\sum_{j=1}^{J} D_{j,t}(Y_{j,t} - Y_{0,t}), 1 - D_{0,0}) \\ &- cov(\sum_{j=1}^{J} D_{j,t-1}(Y_{j,t-1} - Y_{0,t-1}), 1 - D_{0,0}) \ , \end{aligned} \tag{B-2}$$

The first expression in B-2 is zero because of the parallel trend assumption. The second term can be simplified as follows:

$$cov(\sum_{j=1}^{J} D_{j,t}(Y_{j,t} - Y_{0,t}), 1 - D_{0,0}) = -\sum_{j=1}^{J} cov(D_{j,t}(Y_{j,t} - Y_{0,0}), D_{0,0}) \ ,$$

where

$$\begin{aligned} cov(D_{j,t}(Y_{j,t} - Y_{0,t}), D_{0,0}) &= \mathbb{E}[D_{j,t}(Y_{j,t} - Y_{0,t})D_{0,0}] - \mathbb{E}[D_{j,t}(Y_{j,t} - Y_{0,t})]\mathbb{E}[D_{0,0}] \\ &= -\mathbb{E}[Y_{j,t} - Y_{0,t} | D_{j,t} = 1] p_{j,t} p_{0,0} \ . \end{aligned}$$

The second equality follows from the first because $\mathbb{E}[D_{j,t}(Y_{j,t} - Y_{0,t})D_{0,0}] = 0$ for all $j \geq 1$.

The third term in B-2 similarly simplifies to

$$\sum_{j=1}^{J} cov(D_{j,t-1}(Y_{j,t-1} - Y_{0,t-1}), D_{0,0}) = -\sum_{j=1}^{J} \mathbb{E}[Y_{j,t-1} - Y_{0,t-1}|D_{j,t-1} = 1]p_{j,t-1}p_{0,0} \ .$$

Combining all the terms in the numerator with the denominator yields

$$\beta_3^{DiD} = \frac{1}{1 - p_{0,0}} \sum_{j=1}^{J} p_{j,t}\Delta_{j,t}^{TT} - p_{j,t-1}\Delta_{j,t-1}^{TT} \ . \tag{B-3}$$

Proof of Proposition 2.

The following DiD model including time-varying waivers,

$$\begin{aligned} Y_{i,t} =& \beta_0 + \beta_1 Post_t + \beta_2 Treat_i + \beta_3 Post_t \times Treat_i + \\ & \beta_4 W_{i,t} + \beta_5 W_{i,t} \times Post_t + \beta_6 W_{i,t} \times Treat_i + \beta_7 W_{i,t} \times Post_t \times Treat_i + \varepsilon_{i,t} \ . \end{aligned} \tag{B-4}$$

can be written in first differences as

$$Y_{i,t} - Y_{i,t-1} = \beta_1 + \beta_3 Treat_i + (\beta_4 + \beta 5)W_{i,t} + (\beta_6 + \beta_7)Treat_i \times W_{i,t} + \varepsilon_{i,t} - \varepsilon_{i,t-1} \ ,$$

when no waivers have been implemented by the base year of the sample ($W_{i,t-1} = 0$). In this version, β_3 can be estimated by conditioning on the sample with $W_{i,t} = 0$. Similar to the proof of proposition 1,

$$\begin{aligned} \beta_3^{DiD} &= \frac{cov(Y_t - Y_{t-1}, Treat|W_{i,t} = 0)}{Var(Treat|W_{i,t} = 0)} \\ &= \frac{1}{1 - p_{0,t}^{W=0}} \sum_{j=1}^{J} p_{j,t}^{W=0}\Delta_{j,t}^{TT,W=0} - p_{j,t-1}^{W=0}\Delta_{j,t-1}^{TT,W=0} \ , \end{aligned} \tag{B-5}$$

where $p_{j,t}^{W=0} = \mathbb{P}(D_{j,t} = 1|W_{i,t} = 0)$, $\Delta_{j,t}^{W=0} = \mathbb{E}(Y_{j,t} - Y_{0,t}|W_{i,t} = 0)$, $p_{j,t-1}^{W=0}\mathbb{P}(D_{j,t-1} = 1|W_{i,t} = 0)$, and $\Delta_{j,t-1}^{W=0} = \mathbb{E}(Y_{j,t-1} - Y_{0,t-1}|W_{i,t} = 0)$ for $j = 1, \ldots, J$.

Imperfect DiD Interpretation with Additional Restrictions. In the case in which potential outcomes are time-invariant, $Y_{j,t} = Y_{j,t-1} = Y_j$ for $j = 1, \ldots, J$, we can simplify $Y_t - Y_{t-1}$ as follows

$$Y_t - Y_{t-1} = \sum_{j=1}^{J} \sum_{h \neq j} D_{h,t} D_{j,t-1} (Y_j - Y_h) \ .$$

The DiD estimand can then be rewritten as

$$\begin{aligned}
\beta_3^{DiD} &= \frac{cov(Y_t - Y_{t-1}, 1 - D_{0,0})}{var(1 - D_{0,0})} \\
&= \frac{cov(\sum_{j=1}^{J} \sum_{h \neq j} D_{h,t} D_{j,t-1} (Y_j - Y_h), 1 - D_{0,0})}{p_{Treat}(1 - p_{Treat})} \\
&= \frac{1}{p_{Treat}} \sum_{j=1}^{J} \sum_{h \neq j} \phi_{j,h} E[Y_j - Y_h | D_{j,t} = 1, D_{h,t-1} = 1] \ ,
\end{aligned}$$

where $p_{Treat} = \mathbb{P}(Treat = 1)$, and $\phi_{j,h} = \mathbb{P}(D_{j,t} = 1, D_{h,t-1} = 1)$ for $j, h = 1, \ldots, J$.

Imperfect compliance. Suppose that only a subsample of individuals in the treatment group receive any treatment. This is the setting where single motherhood – rather than actual benefit receipt – is defined as the treatment indicator. We define treatment indicators D_j^{τ}, for every EITC benefit $j \in \{0, 1, \ldots J\}$ of treated individuals and retain $D_{0,t} \in \{0, 1\}$ to denote membership in the control group. The treatment indicators are mutually exclusive and collectively exhaustive for each period t so that $D_{0,t} + \sum_{j=0}^{J} D_{j,t}^{\tau} = 1$. We can write the observed outcome as a function of treatment assignments and potential outcomes:

$$Y_t = Y_{0,t} + \sum_{j=0}^{J} D_{j,t}^{\tau} (Y_{j,t}^{\tau} - Y_{0,t}). \tag{B-6}$$

Proposition 3. *Suppose Assumption 1 holds. Then, the DiD estimand in this case is equal to:*

$$\beta_3^{DiD} = \frac{1}{p_{Treat}} \sum_{j=0}^{J} p_{j,t}^{\tau} \Delta_{j,t}^{TT,\tau} - p_{j,t-1}^{\tau} \Delta_{j,t-1}^{TT,\tau} \ , \tag{B-7}$$

where $p_{Treat} = \mathbb{P}(Treat = 1)$, $p_{j,t}^{\tau} = \mathbb{P}(D_{j,t}^{\tau} = 1)$, *and* $\Delta_{j,t}^{TT,\tau} = E[Y_{j,t}^{\tau} - Y_{0,t} | D_{j,t}^{\tau} = 1]$

for $j \in \mathcal{J}$.

Proof. Analogous to proof of Proposition 1. □

Proposition 3 is similar to Proposition 1 in that it highlights that the DiD estimand fails to identify any of the causal parameters of interest without additional restrictions. However, the implications differ if the policy regime did not exist prior to the reform. With imperfect compliance and no pre-existing policy regime, the DiD estimand still does not identify ATT unless we make the additional assumption that the behavior of the treatment group and the control group are identical in the case of no EITC benefits, $Y_{0,t} = Y^{\tau}_{0,t}$. When analyzing the EITC, this means that single mothers and single women without children are assumed to have the same counterfactual outcomes without any tax credits. This restriction is much stronger than Assumption 1 and unlikely to be satisfied, as the literature on female labor supply has documented large differences in labor supply between women and mothers.

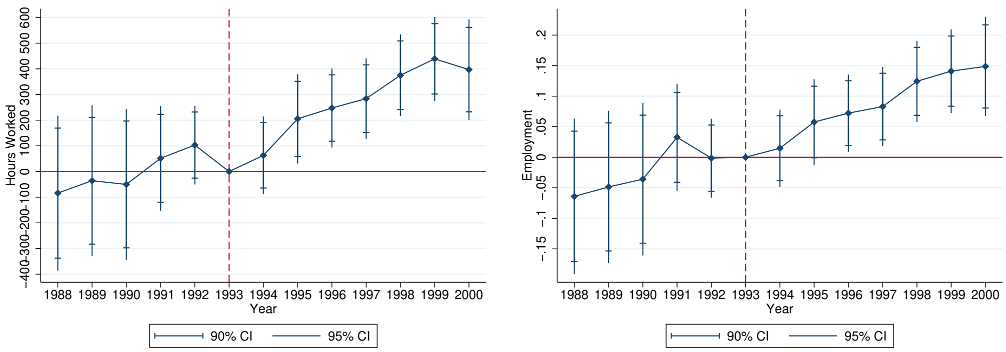
C Additional Tables and Figures

Table C-1: Summary Statistics

Sample of Single Mothers		
	Mean	Standard Deviation
Employment Rate	0.792	
Yearly Hours Worked	1440.430	940.395
Earnings ($ 1000's)	24.552	22.924
EITC Benefits ($ 1000's)	1.008	1.458
Welfare Benefits ($ 1000's)	1.974	3.909
One Child	0.426	
Two Children	0.348	
Three or More Children	0.226	
White	0.661	
Black	0.300	
Other Races	0.039	

This table presents descriptive statistics for the sample of single mothers used in estimation. All monetary values are expressed in thousands of 2015 US dollars.

Figure C-1: Event Study Analysis of the 1993 EITC Reform



(a) Hours: Low-Income Subgroup (b) Employment: Low-Income Subgroup

The figure shows the event study analysis of the effect of the 1993 EITC reform on yearly hours worked (Panel (a)) and employment (Panel (b)) of single mothers. The analysis is run on the subsample of single mothers and single women without children with labor income below \$30,000 (EITC eligibility threshold). The treatment group is made of single mothers and the control group is made of single women without children. Yearly hours worked (Panel (a)) and an indicator variable for employment status (Panel (b)) are regressed on a set of interaction terms between the indicator variable for the treatment group (single mothers) and indicator variables for each year in the period 1988-2000. The event study specification in each panel also contains control variables for the number of dependent children (indicators), state fixed effects, state unemployment level and state welfare waivers (indicator). The year of the reform, 1993, is the reference year for the analysis. The red horizontal line separates the pre-reform (pre-1993) period from the post-reform period. Each panel shows the point estimates for the treatment effect of the reform together with the 90 and 95 percent confidence intervals based on standard errors clustered at the individual level.

Table C-2: EITC Benefits,Welfare Benefits and Weeks Worked per Year

	(1)	(2)	(3)	(4)	(5)	(6)
	Outcome: Weeks Worked per Year					
Change in EITC Benefits ($ 1000s)	3.016***		2.578***	2.688***	2.680***	2.564***
	(0.180)		(0.212)	(0.218)	(0.219)	(0.228)
Change in Welfare Benefits ($ 1000s)		-0.800***	-0.342***	-0.467***	-0.471***	-0.482***
		(0.056)	(0.066)	(0.071)	(0.071)	(0.076)
N	10959	10959	10959	10959	10959	10959
Controls and State F.E.	No	No	No	Yes	Yes	Yes
Unemployment and Waiver Controls	No	No	No	No	Yes	Yes
Control Function	No	No	No	No	No	Yes

The table shows the causal effect of changes in policy-induced EITC benefits and welfare benefits on the change in yearly weeks worked by single mothers. The dependent variable is the year-on-year change in weeks worked by single mothers. Policy-induced changes in EITC and welfare benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC and welfare benefits are provided in Section 3.1 and Appendix D. Control variables include mother's race, indicator variables for the number of dependent children, and year fixed effects. Unemployment and waiver controls include controls for state unemployment level and indicator variables for state welfare waivers. The control function includes a set of controls for lagged labor income, lagged business income, lagged farm income, and lagged non-labor income. See Appendix D for details on the control function. Standard errors are robust to heteroskedasticity and reported in parentheses. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

Table C-3: EITC Benefits,Welfare Benefits and Weeks Worked per Year (by Previous Employment Status)

	(1)	(2)	(3)	(4)	(5)	(6)
	Outcome: Weeks Worked per Year					
Change in EITC Benefits ($ 1000s)	1.94***		1.63***	1.84***	1.83***	1.94***
	(0.57)		(0.58)	(0.60)	(0.60)	(0.60)
Change in Welfare Benefits ($ 1000s)		0.39***	0.34***	0.19*	0.18*	0.17*
		(0.10)	(0.10)	(0.10)	(0.10)	(0.10)
Working in t-1 × Change in EITC Benefits ($ 1000s)	0.33		0.40	0.32	0.32	0.33
	(0.62)		(0.63)	(0.65)	(0.65)	(0.65)
Working in t-1 × Change in Welfare Benefits ($ 1000s)		-1.16***	-0.61***	-0.53***	-0.53***	-0.60***
		(0.15)	(0.16)	(0.17)	(0.17)	(0.17)
N	10959	10959	10959	10959	10959	10959
Controls and State F.E.	No	No	No	Yes	Yes	Yes
Unemployment and Waiver Controls	No	No	No	No	Yes	Yes
Control Function	No	No	No	No	No	Yes

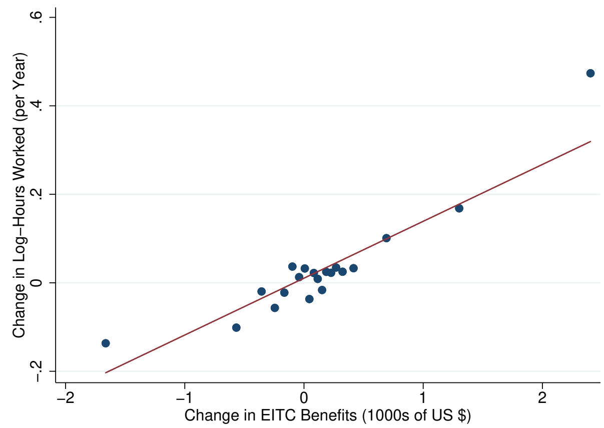
The table shows the causal effect by lagged employment status of changes in policy-induced EITC benefits and welfare benefits on the change in yearly weeks worked by single mothers. The dependent variable is the year-on-year change in weeks worked by single mothers. Policy-induced changes in EITC and welfare benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC and welfare benefits are provided in Section 3.1 and Appendix D. The variable *Working in* $t-1$ is an indicator variable taking the value of one if the mother was working at baseline, and zero otherwise. Control variables include mother's race, indicator variables for the number of dependent children, and year fixed effects. Unemployment and waiver controls include controls for state unemployment level and indicator variables for state welfare waivers. The control function includes a set of controls for lagged labor income, lagged business income, lagged farm income, and lagged non-labor income. See Appendix D for details on the control function. Standard errors are robust to heteroskedasticity and reported in parentheses. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

Table C-4: EITC Benefits,Welfare Benefits and Log-Hours Worked per Year

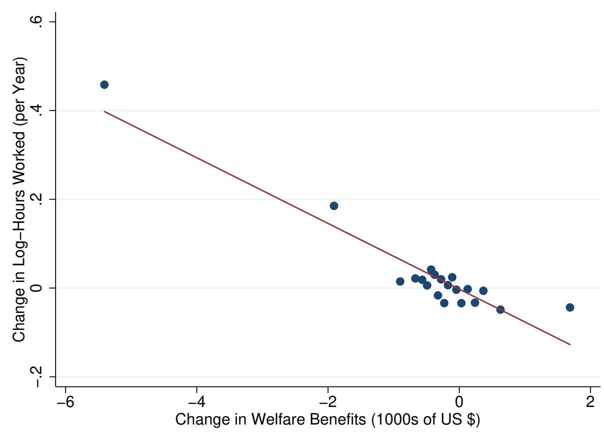
	(1)	(2)	(3)	(4)	(5)	(6)
	Outcome: Yearly Log-Hours Worked					
Change in EITC Benefits ($ 1000s)	0.172***		0.118***	0.124***	0.124***	0.120***
	(0.013)		(0.013)	(0.013)	(0.013)	(0.013)
Change in Welfare Benefits ($ 1000s)		-0.097***	-0.067***	-0.070***	-0.070***	-0.069***
		(0.008)	(0.008)	(0.009)	(0.009)	(0.009)
N	8195	8195	8195	8195	8195	8195
Controls and State F.E.	No	No	No	Yes	Yes	Yes
Unemployment and Waiver Controls	No	No	No	No	Yes	Yes
Control Function	No	No	No	No	No	Yes

The table shows the causal effect of changes in policy-induced EITC benefits and welfare benefits on the change in yearly log-hours worked by single mothers. The dependent variable is the year-on-year change in log-hours worked by single mothers. Policy-induced changes in EITC and welfare benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC and welfare benefits are provided in Section 3.1 and Appendix D. Control variables include mother's race, indicator variables for the number of dependent children, and year fixed effects. Unemployment and waiver controls include controls for state unemployment level and indicator variables for state welfare waivers. The control function includes a set of controls for lagged labor income, lagged business income, lagged farm income, and lagged non-labor income. See Appendix D for details on the control function. Standard errors are robust to heteroskedasticity and reported in parentheses. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

Figure C-2: Log-Hours Worked, Welfare Transfers and EITC Benefits



(a) Log-Hours and EITC



(b) Log-Hours and Welfare

The figure shows the relation between policy-induced changes in EITC benefits (Panel (a)) and welfare benefits (Panel (b)) on the change in yearly log-hours worked by single mothers, thereby conditioning on the set of mothers working positive hours in both periods of the sample. Policy-induced changes in EITC and welfare benefits are expressed in thousands of 2015 US dollars. Details on the construction of the variable for policy-induced changes in EITC and welfare benefits are provided in Section 3.1 and Appendix D. Each panel depicts the binscatter of the nonparametric relation and linear fit line between the y-residuals on x-residuals with specifications containing control variables for mother's race, number of dependent children (indicators), year fixed effects, state fixed effects, state unemployment level and state welfare waivers (indicator).

D Tax and Transfer Rules, and Variables Construction

Construction of independent variables

The challenge in estimating the linear labor supply model in Equation (7),

$$\Delta h_{i,t} = \gamma_0 + \gamma_1 \Delta\xi_{i,t} + \gamma_2 \Delta T_{i,t} + \Delta\epsilon_{i,t} \,,$$

using observed changes in EITC and welfare benefits is that they depend on labor supply. The EITC and welfare influence maternal income in two ways: (i) directly through the transfer, and (ii) indirectly through the labor supply response. This second channel is the source of the endogeneity. Consider an individual with pre-tax earnings, $I_{i,t}^{pre-tax} \equiv \omega_{i,t} \cdot h_{i,t}$, and nonlabor income, $NL_{i,t}$.The benefit formulas for both subsidies depend on labor supply through her pre-tax earnings:

$$\begin{aligned} \xi_{i,t} &= \xi_{i,t}(I_{i,t}^{pre-tax}) = \xi_{i,t}(\omega_{i,t} \cdot h_{i,t}) \,, \\ T_{i,t} &= T_{i,t}(I_{i,t}^{pre-tax}, NL_{i,t}) = T(\omega_{i,t} \cdot h_{i,t}, NL_{i,t}) \,. \end{aligned}$$

To eliminate this source of endogeneity, we calculate policy-induced changes in benefits for each individual caused by variation in the EITC and welfare schedules over time. These policy-induced changes are calculated on the basis of *predicted* earnings and nonlabor income:

$$\Delta\xi_{i,t}\left(I_{i,t-1}^{pre-tax}\right) = \xi_{i,t}\left(\widehat{E}\left[I_{i,t}^{pre-tax} | I_{i,t-1}^{pre-tax}\right]\right) - \xi_{i,t-1}\left(I_{i,t-1}^{pre-tax}\right) \,, \tag{D-1}$$

$$\Delta T_{i,t}\left(I_{i,t-1}^{pre-tax}\right) = T_{i,t}\left(\widehat{E}\left[I_{i,t}^{pre-tax} | I_{i,t-1}^{pre-tax}\right], \widehat{E}\left[NL_{i,t} | NL_{i,t-1}\right]\right) - T_{i,t-1}\left(I_{i,t-1}^{pre-tax}, NL_{i,t-1}\right) \,. \tag{D-2}$$

We follow Dahl and Lochner (2012) and use a fifth-order polynomial in the lagged variable as well as an indicator for a positive lagged value to construct the conditional expectation.

$\Delta\xi_{i,t}$ and $\Delta T_{i,t}$ represent policy-induced changes in the benefits a mother would expect to receive based on first-period income. To the extent that these differ from zero, it is due to factors – such as shifts in policy – that are exogenous with respect to the mother's labor supply decision.

EITC, Welfare, and Food Stamp Formulas

Given $I_{i,t-1}^{pre-tax}$, $NL_{i,t-1}$, and estimates of $\widehat{E}[I_{i,t}^{pre-tax}|I_{i,t-1}^{pre-tax}]$ and $\widehat{E}[NL_{i,t}|NL_{i,t-1}]$, we calculate EITC benefits using NBER's TAXSIM and welfare benefits using the AFDC/TANF rules in effect for each year and state in which we observe the mother. Table A.1 in Kleven (2020) provides a detailed reference for the federal EITC parameters during the period we study.

The computation of welfare benefits depends on earnings, nonlabor income, number of children, and the individual's state of residence. Each state bases eligibility on whether both gross and net income fall below a threshold specific to the number of children in the family. If a family is eligible, they receive a benefit that depends on several parameters set by the state: the maximum allowable benefit (MB), the dollar and rate disregards to earnings (EDD and ERD) described in Appendix A, and the payment standard (PS). Given earnings ($I_{i,t}^{pre-tax}$) and nonlabor income ($NL_{i,t}$), net income is given by $Net_{i,t} = (I_{i,t}^{pre-tax} - EDD)(1 - ERD) + NL_{i,t}$, and benefits are

$$Benefit_{i,t} = \max\{\min\{MB, PS - Net_{i,t}\}, 0\} \ . \tag{D-3}$$

All parameters vary substantially across states and years. Figure A-2 plots the welfare benefit function for four states in selected years between 1988 and 2002 for a mother with 2 two children and no nonlabor income. MB determines the y-intercept, while ERD influences the slope of the function and a positive EDD induces a benefits to be constant in earnings for low levels of earnings.

SNAP benefits (food stamps) enter into the individual's budget constraint in the model in Section 4. SNAP benefits depend on an individual's earnings, nonlabor income, and welfare benefits received. Provided that gross earnings are below a 130 percent of the federal poverty line and earnings and nonlabor income net of welfare benefits ($NE_{i,t}$) are below 100 percent of the poverty line, an individual with k children receives SNAP benefits according to the formula:

$$SNAP_{i,t}(k) = \max\{MB_{i,t}(k) - 0.3 * NE_{i,t}(k), 0\} \ , \tag{D-4}$$

where $MB_{i,t}(k)$ is the maximum benefit for a family with k children in year t.

Control Function

$\Delta\xi_{i,t}$ and $\Delta T_{i,t}$ are functions solely of $I_{i,t-1}^{pre-tax}$ and $NL_{i,t-1}$. To account for these possible correlations, we introduce control functions in these variables, $\phi_1(I_{i,t-1}^{pre-tax})$ and $\phi_2(NL_{i,t-1})$ in the specification in column (6) of Tables 1 - 2. We use lagged values of labor income, non-labor income, business income, and farm income as part of the control function. The empirical results show that our findings are not sensitive to the inclusion of the control functions.

Parameters of Estimated Tax Function

We approximate mother's after-tax income by the parametric function in 11. For each year, state, and number of children, we use NBER's TAXSIM program to simulate the aftertax earnings of mothers with incomes at intervals of $1000 between $0 and $100,000. Then we estimate $\theta_{0,s,t,k}$ and $\theta_{1,s,t,k}$ $\forall s, t, k$ by minimizing the sum of squared residuals between actual after-tax income and the after-tax income predicted by the right-hand-side of Equation (11). Estimation is done by Nonlinear Least Squares.

E Welfare Time Limits

The table in this section documents the earliest possible date on which time limits might result in a welfare recipient being kicked off the welfare rolls (US Department of Health and Human Services 1997).

State	**Extent**	**First Cases Reach Limit**	**Consequence**
Arizona	Statewide	November-97	Adult portion of grant is terminated.
California	Statewide	August-97	Adults must participate in CWEP for 100 hours per month.
Colorado	Five counties	May-96	Non-exempt adults must be working at least 30 hours per week or actively participating in a JOBS training program.
Connecticut	Two cities: New Haven and Manchester	June-96	End of cash assistance.
Connecticut	Statewide	September-97	End of cash assistance.
Delaware	Statewide	November-97	Adult must enter pay-after-performance work experience program.
Delaware	Statewide	November-99	End of cash assistance.

Florida	Escambia & Alachua counties. Later expanded to six more counties.	February-96	End of cash assistance. Transitional employment will be offered to for those who have diligently completed plans, are unable to find employment and have not voluntarily quit or been discharged for misconduct.
Georgia	Ten counties	Between December-98 and December-99	Recipients must work 20 hours per month in a work experience program for a state, local government, federal agency or nonprofit organization, subject to availability of work slots.
Hawaii	Statewide	Between November-01 and November-02	End of cash assistance.
Illinois	Statewide	November-96	Recipients whose youngest child is 13 or older must accept up to 60 hours per month of work subsidized by AFDC grant.
Illinois	Statewide	November-97	End of cash assistance; family ineligible to reapply for aid for two years.

Indiana	Statewide (initially limited to 12,000 adult recipients)	May-97	Adult portion of grant is terminated.
Iowa	Statewide	Unknown	Benefits will be phased out for failure to make satisfactory progress towards self-sufficiency.
Louisiana	Statewide	February-99	End of cash assistance.
Massachusetts	Statewide	January-96	Recipients who can not find work will be placed in a community service position for 20 hours per week.
Michigan	Statewide	November-95	After one year of non-compliance with work requirements, penalty increases to loss of all AFDC benefits.
Missouri	Statewide	June-97	At the time limit, recipients will be assigned to job search and work experience.

Missouri	Statewide	May-98	The state will deny AFDC to an individual who received benefits for at least 36 months and who reapplies after completing a self-sufficiency agreement entered into after July 1, 1997, if the individual was responsible for becoming unemployed. Other eligible members of the family will receive benefits.
Montana	Statewide	February-98 for single parents	Individuals who reach time limit but have not achieved self-sufficiency will be required to participate in Community Services Program for 20 hours per week in order to receive benefits.
Nebraska	Two counties in 1995. Expanded statewide in 1996.	November-97	End of cash assistance.
New Hampshire	Statewide	Between February-97 and February-98	Requires job search for up to 26 weeks followed by work-related activities for 26 weeks. These cycles will repeat until the recipient is off AFDC.

North Carolina	Statewide	Between March-98 and March-99	End of cash assistance. Family becomes ineligible for 36 months.
North Dakota	Ten counties	Unknown	Placement in a work experience program or extension of benefits, based on an evaluation of the recipient's circumstances.
Ohio	Statewide	Between July-99 and January-00.	End of cash assistance.
Oklahoma	Six counties	May-99	Mandatory workfare participation of at least 24 hours a week.
Oregon	Statewide	August-97	End of cash assistance.
South Carolina	Statewide	Between July-98 and July-99	End of cash assistance.
South Dakota	Statewide	May-96 or May-99	If adult is not employed at least 30 hours per weeks, must perform 30 hours of approved volunteer service each week (fewer if good cause shown).
Tennessee	Statewide	Between April-98 and April-99 for continuous recipients	End of cash assistance. After receiving AFDC for 18 months, a household must wait at least three months before re-applying.

Texas	Statewide	May-97	Adults who reach the time limits may not receive cash assistance for a five-year period. The children will continue to be eligible for benefits.
Vermont	Statewide	November-95 for 15-month group. March-97 for 30-month group.	Requires participation in subsidized employment.
Virginia	Statewide. Phased in over four years.	August-97	End of cash assistance.
Washington	Statewide	February-00	Imposes a 10 percent grant reduction for families who have received assistance for 48 out of 60 months, and imposes an additional 10 percent grant reduction for every 12 months thereafter.
Wisconsin	Two counties: Fond du Lac and Pierce	February-97	End of cash assistance.